Nonlinear exchange rate pass-through in timber products: the case of oriented strand board in Canada and the United States

Barry K. Goodwin and Matthew T. Holt and Jeffrey P. Prestemon

University of Alabama, North Carolina State University, USDA Forest Service

22 August 2012

Nonlinear Exchange Rate Pass–Through in Timber Products: The Case of Oriented Strand Board in Canada and the United States

Barry K. Goodwin*
Department of Agricultural and Resource Economics
North Carolina State University

Matthew T. Holt†
Department of Economics, Finance, and Legal Studies
University of Alabama

Jeffrey P. Prestemon‡
USDA Forest Service
Southern Research Station

This Draft: August 22, 2012

Abstract

We assess exchange rate pass–through (ERPT) for U.S. and Canadian prices for oriented strand board (OSB), a wood panel product used extensively in U.S. residential construction. Because of its prominence in construction and international trade, OSB markets are likely sensitive to general economic conditions. In keeping with recent research (e.g., Al-Abri and Goodwin, 2009; Larue et al., 2010), we examine regime–specific ERPT effects; we use a smooth transition vector error correction model. We also build on work by Nogueira, Jr. and León-Ledesma (2011) and Chew et al. (2011) in considering ERPT asymmetries associated with a measure of general macroeconomic activity. Our results indicate that during expansionary periods ERPT is modest, at least initially, but during the recent financial crises ERPT effects were quite large.

Keywords: Exchange rate pass–through, oriented strand board, smooth transition model, unemployment

JEL Classification Codes: E32, F10; F30; F41; L16

*Department of Agricultural and Resource Economics and Department of Economics, North Carolina State University, Campus Box 8109, Raleigh, NC 27695–8109, USA. Telephone: 919-515-4620. Fax: 765-515-1824. E-mail: barry_goodwin@ncsu.edu.

†**Corresponding Author**: Department of Economics, Finance, and Legal Studies, University of Alabama, Box 870224, Tuscaloosa, AL 35487-0224, USA. Telephone: 205-348-8980. Fax: 205-348-0590. E-mail: mtholt@cba.ua.edu.

‡Forestry Sciences Laboratory, P.O. Box 12254, 3041 Cornwallis Road, Research Triangle Park, NC, 27709-2254, USA. Telephone: 919-549-4033. Fax: 919-549-4047. E-mail: jprestemon@fs.fed.us.

1 Introduction

Questions regarding the extent of exchange rate pass–through (ERPT) into import prices, in other words, the degree to which exchange rate shocks evoke equilibrating price response for traded commodities and goods, have long been of interest to economists and policy makers. Much of the recent interest in this topic can perhaps be traced to the observation that estimated ERPT effects are generally reported to be small (Goldberg and Knetter, 1997) For example, a widely cited rate of pass–through into aggregate import price is approximately 50%, as reported by Goldberg and Knetter (1997). In addition to relatively low rates of ERPT, there is also mounting evidence that they have been declining over time; see, for example, Bailliu and Fujii (2004), Campa and Goldberg (2005), and Marazzi and Sheets (2007), among others. Correspondingly, several strands of the ERPT literature have evolved. One is the so called macro strand, where the focus is on determining the extent of ERPT to import prices at the aggregate level and, secondarily, the extent to which such responses are passed along to consumers (see, e.g., Gagnon and Ihrig, 2004). Another strand focuses on determining the extent to which ERPT impacts import prices at the industry or commodity level, where incomplete pass–through is often conjectured to be a function of the market structure of the industry being examined. Examples of work in this vain include Knetter (1989) and Pollard and Coughlin (2004). Of interest is that empirical estimates of long–run ERPT at the industry or commodity level are often even smaller than those obtained by using more aggregated data.

Over the years various theories and/or methodological refinements have been explored in an attempt to account for low and/or declining rates of ERPT. Of interest is that a small number of recent studies have examined the possibility that there are asymmetries or nonlinearities in pass–through, that is, for example, that a currency depreciation could have different impacts on import prices than would an appreciation or, similarly, that large changed may have different effects than small ones. In one of the earliest studies of this sort, Mann (1986) found evidence of asymmetric pass–through effects. Likewise, by employing aggregate data for seven Asian Pacific countries, Webber (2000) reports substantial evidence of asymmetric pass–through effects for five of these. Bussiere (2007) considers pass–through into import and export prices in G7 countries, and finds substantial evidence of nonlinearities. Even so, Bussiere (2007) only tests for nonlinearity and does not otherwise estimate corresponding nonlinear models of pass through. Karoro, Aziakpono and Cattaneo (2009) consider asymmetries in pass–through to import prices in South Africa; they find evidence that ERPT is higher during periods of rapid appreciation relative to deprecation. As well, Al-Abri and Goodwin (2009) update the data used by Campa and Goldberg (2005) and also allow for threshold effects with respect to ERPT into G7 country import prices. Overall, they find substantial evidence of non-linearities in pass–through effects. In a closely related study, Larue, Gervais and Rancourt (2010) examine the possibility asymmetric ERPT into export prices for pork meat from Canada to Japan and the U.S. by using threshold cointegration techniques.

Many of the studies outlined above have focused on estimating pass–through effects by using either import prices at either the aggregate level or for specific industries. Comparatively few studies have focused on pass–through effects at the individual commodity level. In part this is because commodities are typically homogeneous and to be traded in something close to perfectly competitive market conditions. The implication is that the ability of exporting firms to exert any market power over pricing combined with the perfect arbitrage conditions of the "law of one price" (LOP) are thought to result in complete ERPT for commodity import prices. In short, commodities are thought to have flexible or flex import prices. Even so, there is evidence that, at least in some instances, there is incomplete pass–through for commodities. Jabara and Schwartz (1987) explore ERPT for Japanese import prices for five agricultural commodities, and find evidence of incomplete pass–through as well as evidence of asymmetric responses to exchange rate shocks for several commodities. As well, they find substantial evidence of asymmetric responses to exchange rates for several commodities. Likewise, Uusivuori and Buongiorno (1991) examine ERPT for a number of U.S. forest product exports to Europe and Japan, and find both that pass–through is incomplete and that its effects are asymmetric depending on whether the exchange rate is appreciating or depreciating. Finally, Parsley (1995) examines ERPT for five specific products exported from Japan to the United States. In this study asymmetry in (real) exchange rate effects were also allowed for; the results show there is are apparent declines in ERPT during periods of dollar appreciation.

In general ERPT is an important indicator of the operation and performance of markets for internationally–traded commodities such as OSB. A lack of pass–through may reflect imperfect arbitrage, inefficient trade, inflexible prices (perhaps due to contracts or menu pricing practices), price discrimination, high transactions costs, and the influences of government policies. A lack of full pass–through indicates that standard arbitrage behavior, which is often assumed to hold in absolute terms in conceptual and empirical models of trade, may in fact not be supported empirically. In any event, attaining deeper insights into the nature of ERPT at the primary commodity level is an important agenda in the modern empirical trade literature; there is scope for further work.

To begin, it is surprising that comparatively few studies have explored ERPT at the product or commodity level. As well, while there is mounting evidence that asymmetries or, more generally, nonlinearities are a feature of the exchange rate effect on import prices, it is also surprising that comparatively few studies have examined these effects by using modern time series methods, and especially so when ERPT is examined at the commodity level.[1] The overall goals of this paper are then: (1) to examine ERPT in import prices for a highly traded, homogeneous commodity; and (2) to examine in a general testing and estimation framework the role of nonlinearities in ERPT. Specifically, we examine the (potentially nonlinear) impacts of exchange rates on U.S. import prices and Canadian export prices for oriented strand board (OSB). Oriented strand board represents an interesting case study for which to examine ERPT at the product level. It is a homogeneous product that is widely used in residential and commercial construction throughout North America. As illustrated in Figure 1, in recent years the U.S. has produced more OSB than Canada, but Canada exports both a far higher amount as well as a greater percentage of its total production than does the United States (on average 84% versus 1.6%). Moreover, as also illustrated in Figure 1 the overwhelming majority of all Canadian OSB exports are destined for the United States. While prior work has examined pass–through issues for international trade in various timber products (see, e.g., Uusivuori and Buongiorno, 1991; Bolkesjø and Buongiorno, 2006), to our knowledge similar questions have not been addressed for panel products manufactured wood products. Taken together the evidence suggests that additional insights into ERPT at the product level can be attained by conducting a careful analysis of U.S. and Canadian OSB price relationships.

[1]Notable exceptions include, of course, Al-Abri and Goodwin (2009) and Larue, Gervais and Rancourt (2010), who do use threshold cointegration methods to estimate asymmetric pass–through effects.

2 Conceptual Framework

There is a vast literature that examines questions regarding the law of one price in the context of international (regional) price behavior; see, for example, Goodwin, Holt and Prestemon (2011) along with references therein for a recent review of this research. The micro–foundations underlying exchange rate pass–through are identical to those that motivate the LOP; however, investigations of ERPT highlight the separate effects of price and exchange rate shocks in commodities that are traded across markets with different currencies. In that it is common for internationally–traded commodities to be invoiced in a common currency across different national markets (e.g., the U.S. dollar or the Euro), exchange rates may still have an impact on price linkages if the internal markets being considered have different currencies.

Following Goldberg and Knetter (1997), in the classical pass–through literature the basic long–run price relationship may be stated as:

$$P_{it} \;=\; E_t^{\beta_1} P_{xt}^{\beta_2},\;\; \beta_1,\;\; \beta_2 > 0, \quad t = 1, \ldots, T, \tag{1}$$

where P_{it} is the (nominal) import price in country i for the good in question in period t (denominated in country i's currency); P_{xt} is the corresponding (nominal) export price in country j (denominated in country j's currency); and E_t is the nominal exchange rate, expressed in terms of the importer's currency (i.e., country i's) relative to the exporter's currency (i.e., country j's). As well, β_1 and β_2 are parameters such that with perfect pass–through $\beta_1 = \beta_2 = 1$. It is natural to convert (1) to natural logarithmic form, so that the price relationship may be written as:

$$p_{it} \;=\; \beta_1 e_t + \beta_2 p_{xt}, \tag{2}$$

where lower case letters denote variables expressed in natural log form. If (2) is estimated as is (and for the moment ignoring any possible time series complications associated with the data), then a model for testing the impact of exchange rates on import prices could be specified simply as:

$$p_{it} \;=\; \beta_1 e_t + \beta_2 p_{xt} + \varepsilon_t, \tag{3}$$

where β_1 and β_2 are parameters to be estimated and ε_t is an additive error term such that $\varepsilon_t \sim iid(0, \sigma^2)$. In this case a test of full (complete) exchange rate pass–through would be associated with a test of the hypothesis $\mathrm{H}_0 : \beta_1 = \beta_2 = 1$.

The specification defined by (1)–(3) assumes p_{xt} is measured in the exporter's currency. In the case where exports are invoiced in the importer's currency, the exporter's price may be written as $P_{xt} = \tilde{P}_{xt}/E_t$, where $\tilde{P}_{xt}$ is the export price expressed in the importing country's currency. In this later case, that is, when prices are invoiced in the importer's currency, the model in (3) may be rewritten as:

$$p_{it} \;=\; (\beta_1 - \beta_2)\, e_t + \beta_2 \tilde{p}_{xt} + \varepsilon_t. \tag{4}$$

For complete pass–through we again require $\beta_1 = \beta_2 = 1$, which in turn reduces (4) to a stochastic version of the law–of–one–price relationship. In other words, with common currency pricing complete pass–through implies that exchange rates should have no long–term (permanent) impact on the import price.

Following recent literature (see, e.g., Campa, Goldberg and Gonzalez-Minguez, 2005), we could further modify the model to allow for the possibility that exporting firms, presumably operating in an imperfectly competitive market environment, could maintain a fixed percentage markup over their marginal cost. The assumption of imperfectly competitive market conditions seems relevant for North American OSB markets. In 2006, a series of lawsuits were consolidated into a single case in the U.S. District Court in Pennsylvania on behalf of aggrieved parties involved in OSB purchases between June, 2002 and February, 2006. The suite alleged that the a number of major North American OSB manufacturers, operating in both the United States and Canada, conspired to maintain artificially high prices for OSB during the June, 2002 through February, 2006 period.[2] In any case, to modify the model to allow for imperfectly competitive behavior, we can rewrite $\tilde{p}_{x,t}$ as:

$$\tilde{p}_{xt} \;=\; mkup_{xt}(e_t) + mc_{xt}, \tag{5}$$

where $mkup_{xt}(e_t)$ denotes the percentage markup and mc_{xt} denotes marginal cost, both in logarithmic form. As well, and as indicated by the notation in (5), the markup may also vary with the

[2] A settlement between plaintiffs and defendants was reached in 2008, and subsequently approved by the court in December, 2008. The cases against OSB manufacturers were subsequently dismissed.

exchange rate. Again, following Campa, Goldberg and Gonzalez-Minguez (2005), we may write the markup in (5) as:

$$mkup_{xt}(e_t) \; = \; \phi + \Phi e_t, \tag{6}$$

where ϕ is a component of the markup that does not change with the exchange rate.

In the case were import prices are invoiced in importing firm's currency (i.e., local currency pricing), we may substitute (5) and (6) into (4) to obtain:

$$p_{it} = \alpha_0 + (\beta_1 + \beta_2 (\Phi - 1)) e_t + \beta_2 mc_{xt} + \varepsilon_t, \tag{7}$$

where $\alpha_0 = \phi\beta_2$. Several important observations may be drawn from (7). To begin, even if $\beta_1 = \beta_2 = 1$ holds, incomplete pass–through may occur to the extent that exporting firms operate in an imperfectly competitive market environment. Secondly, if (7) is viewed as a long–run relationship, then we might still reasonably expect a non–zero intercept term to be present if, in fact, ϕ, the constant mark–up parameter, is non–zero.[3] In the literature there are many example of variants of (7) being used to estimate ERPT effects. See, for example, Campa and Goldberg (2005).

The basic framework outlined above can be modified if there is reason to suspect that ERPT is regime specific, that is, that the impact of exchange rates on rates on import prices varies with either the magnitude or direction of adjustment of some other variable including but not limited to the exchange rate itself. To illustrate, as Al-Abri and Goodwin (2009) and Larue, Gervais and Rancourt (2010) note, the markup equation in (6) might be such that the exchange rate response parameter, Φ, varies depending on the size (or sign) of a change in exchange rates. For relatively small exchange rate adjustments exporters may decide not to adjust the markup due to menu costs. But for a large exchange rate adjustment, exporting firms may be forced to adjust markups in order to maintain market share. Alternatively, under local currency pricing the exporting firm presumably must still convert revenues earned in foreign currency into the home currency. Presumably doing so involves transactions costs and, moreover, costs that might vary with the magnitude of recent exchange rate movements.

[3] In addition, α_0 may also capture factors associated with the cost of trade if such factors are proportional to prices, an assumption that is in turn common empirical studies of price parity relationships.

In any event, the model in (7) can be modified in the following manner:

$$p_{it} = \alpha_0 + \left(\beta_1 + \beta_2 \left(\Phi_1 \left(1 - I_{\{s_t > \theta\}}\right) + \Phi_2 I_{\{s_t > \theta\}} - 1\right)\right) e_t + \beta_2 mc_{xt} + \varepsilon_t, \tag{8}$$

where θ is the threshold parameter, and where $I_{\{s_t>\theta\}}$ is a Heaviside indicator function such that $I_{\{s_t\leq\theta\}} = 1$ if $s_t > \theta$ and is 0 otherwise. Here s_t is the so called transition variable; it is the variable that, in conjunction θ, determines the nature of nonlinear pass–through effects. Let $s_t = f(z_t)$, where z_t is some underlying variable and the form of the function $f(.)$ is presumably known. For example, and as already noted, z_t might equal e_{t-1}, the (lagged) exchange rate, although z_t could be equated with other observed variables as well. Regarding the specification of s_t in (8), it might be that: (1) $s_t = z_{t-k}$; or (2) that $s_t = z_{t-1} - z_{t-2}$; or (3) that $s_t = (z_{t-1} - z_{t-2})^2$. See, for example, van Dijk, Teräsvirta and Franses (2002) for additional details. In any event, the important point is that the markup varies depending on recent movements in z_t, and therefore ERPT effects will also vary with these changes.

As already noted, nonlinearities in ERPT effects could arise for reasons other than those associated with exchange rate behavior. For example, in a recent study Chew, Ouliaris and Tan (2011) considered import prices for Singapore, and allowed the ERPT effects into these prices to vary with the business cycle.[4] Their results confirmed there are asymmetric pass–through effects into Singapore's import price over the business cycle, with smaller pass–through occurring during expansions as compared with retractions.

Business cycle effects with respect to ERPT in North American OSB markets might be especially relevant given that residential construction (a primary end–use for OSB) is quite sensitive to economic downturns; indeed, housing starts are often asserted to be an important leading indicator of overall economic activity (Leamer, 2007). As an empirical proposition then, it is certainly plausible that markups and hence ERPT could vary with the business cycle even when considering price response for a specific commodity such as OSB. In terms of (8), the idea would be to link z_t and hence s_t to one or more variables that transmit information regarding the stage of the business cycle.

[4] Chew, Ouliaris and Tan (2011) accomplish this by using a band–pass spectral regression of the long–run pass–through relationship. The basic idea is that the equation's parameters vary with the phase of the business cycle, thereby allowing for nonlinear ERPT effects in the model.

3 Data

3.1 Data Description

As indicated previously, we focus on prices for oriented strand board (OSB) in Canada and the United States. OSB is a manufactured wood product that was introduced in 1978, and is widely used in residential and commercial construction, with the bulk of OSB produced in North America originating in the Southern U.S. and Canada. For example, in 2009 and 2010 Canada and the Southern U.S. produced nearly ninety–percent of all OSB otherwise produced in North America (Engineered Wood Product Association, 2010). OSB is constructed by using waterproof and heat cured resins and waxes, and consists of rectangular shaped wood strands that are arranged in oriented layers. As well, it is manufactured in long, continuous mats which are then cut into panels of varying sizes. As a panel product OSB is similar to plywood, although it is generally considered to have more consistency than plywood and is cheaper to produce. As indicated in Figure 1, the Structural Board Association (SBA) reports that in 1980 OSB panel production in the U.S. was 135 million square feet (on a $3/8^{th}$'s inch basis) and in Canada was 616 million square feet. Comparable numbers for 2010 were 10,838 million square feet produced in the U.S. and 4,700 million square feet in Canada. The SBA also reports that by 2000 OSB production exceeded that of plywood, and that by 2010 OSB production enjoyed a 58–percent market share among all panel products in North America. Figure 1 illustrates the substantial growth in OSB production since 1995 as well as the sharp decline in OSB production following the collapse of the U.S. housing market in 2007.

Considering the above, we focus on pass–through effects for OSB in two regional North American markets: (1) Eastern Canada (production deriving from plants in Ontario and Quebec); and (2) the Southeast U.S. (production deriving from plants in Georgia, Alabama, Mississippi, South Carolina, and Tennessee). The price data are for panels of $7/16^{th}$'s inch oriented strand board, and are expressed in U.S. dollars per thousand square feet, that is, Canadian mills engage in local currency pricing. All price data are observed on a weekly basis and were obtained from the industry source *Random Lengths*.[5] The regional OSB price data used are FOB mill price averages. The

[5] *Random Lengths* is an independent, privately owned price reporting service, providing information on commonly produced and consumed wood products in the U.S., Canada, and other countries since 1944. Reported open–market sales prices are based on hundreds of weekly telephone interviews with producers, wholesalers, distributors, secondary manufacturers, buying groups, treaters, and large retailers.

period covered is from October 9, 1998 through August 20, 2010, the result being there are 620 usable weekly observations. A plot of the regional OSB price data converted to natural log form is reported in Figure 2. In the analysis we propose treating the (natural logarithm) of the Southeast U.S. as the effective import price (p_i) and, following Wickremasinghe and Silvapulle (2004) and Karoro, Aziakpono and Cattaneo (2009), using the observed (natural logarithm) of the FOB mill price in Eastern Canada (p_x) as a proxy for the exporter's price (marginal cost) in (7) or, respectively, (8). Doing so is reasonable in part because, although the bulk of OSB in the U.S. is produced in the Southeast, it is also the region with the largest growth in demand–Census Bureau data on housing starts confirm that states in the Southeast have, since the late 1980s, dominated much of the rest of the country in terms of overall starts as well as growth in new home construction.

Aside from reasonable proxies for OSB import and export prices, the specification in equation (7) indicates that a relevant exchange rate is also needed. Here we use the (reciprocal of) the week–ending average of the Canadian Dollar–to–U.S. dollar exchange rate as reported on the St. Louis Federal Reserve's Federal Reserve Economic Data (FRED) archive. A plot of the (natural logarithm) of the weekly exchange rate, (e), over the sample period, that is, over the October 9, 1998 through August 20, 2010 period, is also recorded in Figure 2. As illustrated there, the U.S. dollar tended to appreciate relative to the Canadian dollar during the sample period.

Internationally traded commodities such as OSB are likely to be sensitive to economic conditions in the aggregate economy. In the case of OSB, a principal building material used in residential and commercial construction, this is especially likely to be true. To allow for the possibility that changes in the overall economy may affect linkages and exchange rate relationships for U.S. and Canadian OSB markets, an indicator of weekly changes in overall economic conditions is needed. In our case, we use the most frequently cited indicator of the overall health of the economy–the unemployment rate. In particular, we consider weekly, end–of–period insured unemployment claims.[6] These measures are regarded as a reliable indicator of real, aggregate economic activity (Stock and Watson, 2003). Weekly unemployment claims are collected by the U.S. Department of Labor, and are reported on the St. Louis Federal Reserve's FRED online database. The unemployment measure used here, *une*, is the percentage unemployment claims variable without seasonal adjustment. A

[6]Alternatively, the indicator might be linked more directly to some measure of housing starts (Leamer, 2007). The U.S. Census Bureau reports housing starts data, but unfortunately they are at most available on a monthly basis. For this reason we do not pursue this option further in the present study.

plot of the unemployment variable over the sample period is reported in Figure 3.

3.2 Data: Preliminary Properties

Having identified the series to be used in the empirical analysis, it is useful to examine some of their basic statistical properties. Specifically, we test each series for the null of a unit root by using augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests (Dickey and Fuller, 1979; Phillips and Perron, 1988). In implementing the ADF test, we account for the potential effects of heteroskedasticity by using the modified test statistic suggested by Demetrescu (2010). As well, we choose lag lengths for the autoregressive parameters in the ADF test by using the lag–length selection procedures outlined by Ng and Perron (1995); for the PP test we choose a lag length based on the rule $int(4(T/100)^{0.25}$, which is six in the present case. The results are reported in upper panel of Table 1.

As recorded in the Table, the tests provide evidence of nonstationarity for each variable considered.[7] In terms of the conceptual framework outlined in the previous section, the implication is that equation (7) should now be viewed as a cointegrating regression, and thereby reflects the long–run relationship the two price variables and the exchange rate variable. Following Balke and Fomby (1997), we estimate the (unrestricted) version of (7) and test the resulting residual series for the presence of a unit root. The results in this instance are reported in the lower panel of Table 1. Regardless of which test is employed (i.e., ADF or PP), it is clear that we reject the unit root hypothesis and conclude that the prices and the exchange rate are cointegrated. This information will be fundamental in specifying and estimating the subsequent nonlinear model used to estimate ERPT effects, to which we now turn.

4 Modeling Framework

4.1 Multivariate Smooth Transition Models

To explore the the exchange rate pass–through effects for Canadian and U.S. prices, we follow prior literature in specifying a (nonlinear) vector error correction model (VECM) (see, e.g., Al-Abri and Goodwin, 2009). Specifically, the basic building block of our empirical analysis is a

[7] We note, however, that the test statistics for the Eastern Canada OSB price are close to the five–percent critical values.

VECM model of the general form:

$$\Delta \boldsymbol{y}_t = \boldsymbol{\delta} + \sum_{i=1}^{p-1} \boldsymbol{\Psi}_i \Delta \boldsymbol{y}_{t-i} + \boldsymbol{\alpha}\hat{\varepsilon}_{t-1} + \boldsymbol{v}_t, \tag{9}$$

where $\boldsymbol{y}_t = (p_{it}, p_{xt}, e_t)'$; $\hat{\varepsilon}_{t-1}$ is the lagged residual from the cointegrating regression described in the previous section, that is, the (lagged) departure from long–run equilibrium; $\boldsymbol{\delta}$ and $\boldsymbol{\alpha}$ are conformable parameter vectors, where $\boldsymbol{\alpha}$ contains the so called speed–of–adjustment parameters or error correction coefficients; $\boldsymbol{\Psi}_i$ are conformable parameter matrices; and $\boldsymbol{v}_t$ is a vector of mean zero, random, additive errors.

If nonlinear ERPT effects are not considered, then the system in (9) can be estimated and impulse response functions generated in order to determine the degree or pass–through. Alternatively, if nonlinearities of the sort described in previous sections are considered, then it is necessary to modify (9). In the spirit of the regime switching framework in (8), we could re–specify the VECM as:

$$\begin{aligned} \Delta \boldsymbol{y}_t &= \left[\boldsymbol{\delta}_1 + \sum_{i=1}^{p-1} \boldsymbol{\Psi}_{i1} \Delta \boldsymbol{y}_{t-i} + \boldsymbol{\alpha}_1\hat{\varepsilon}_{t-1}\right](1 - G(s_t, \boldsymbol{\theta})) \\ &+ \left[\boldsymbol{\delta}_2 + \sum_{i=1}^{p-1} \boldsymbol{\Psi}_{i2} \Delta \boldsymbol{y}_{t-i} + \boldsymbol{\alpha}_2\hat{\varepsilon}_{t-1}\right] G(s_t, \boldsymbol{\theta}) + \boldsymbol{v}_t, \end{aligned} \tag{10}$$

where Δ is a difference operator such that $\Delta x_t = x_t - x`_{t-1}$. In (10) the function $G(.)$, the so called transition function, now plays the role of the Heaviside indicator function defined previously and $\boldsymbol{\theta}$ is now a vector of parameters that identifies the transition function. Importantly, similar to the Heaviside indicator function, the function $G(.)$ is bounded between zero and one. A primary difference, however, is that $G(.)$ can also assume intermediate values on the unit interval, that is, regime change can be potentially gradual or smooth. For this reason the model in (10) is referred to as a smooth transition VECM, or STVECM, and was introduced originally by Rothman, van Dijk and Franses (2001). Furthermore, the STVECM is a straightforward extension of the univariate smooth transition autoregressive (STAR) models introduced originally by Teräsvirta (1994). The model is, of course, nonlinear in parameters given that γ and c must also be estimated, and therefore nonlinear estimation methods must be employed.

To implement the STVECM it is necessary to specify a form for the transition function, $G(.)$. In the present case if, for example, it is hypothesized that ERPT varies with the magnitude of the departure from long–run equilibrium, then it would be feasible to specify the transition function as:

$$G\left(s_t;\gamma,c\right)=1-\exp\left(-\gamma\left(s_t-c\right)^2\Big/\hat{\sigma}^2_{st}\right), \tag{11}$$

that is, the exponential distribution, where $\boldsymbol{\theta}=(\gamma,c)$, γ being the speed–of–adjustment parameter and c being the centrality parameter; and $\hat{\sigma}_{s_t}$ is the sample standard deviation of the transition variable, s_t. In (11) as the transition variable, s_t, approaches c the function $G(.)$ approaches zero while, conversely, when s_t deviates far from c the function $G(.)$ approaches unity. The speed with which the transition from one extreme to the other occurs is dictated by the magnitude of the parameter, γ. In this manner the exponential function is capable of approximating something akin to a three–regime threshold model of the sort employed by Al-Abri and Goodwin (2009) and Larue, Gervais and Rancourt (2010), albeit in a potentially smooth way. To abbreviate, we refer to a regression equation with an exponential transition function as an exponential smooth transition regression equation, or ESTR.

Alternatively, the logistic function, specified as:

$$G\left(s_t;\gamma,c\right)=\left[1+\exp\left(-\gamma\left(s_t-c\right)/\hat{\sigma}\right)\right]^{-1}, \tag{12}$$

is an another widely used specification for the transition function, $G(.)$, in the STVECM (see, e.g., Rothman, van Dijk and Franses, 2001). In (12) as s_t increases above the centrality parameter c, the function $G(.)$ will approach unity. Alternatively, for s_t below c the logistic function approaches zero. Again, the speed with which this transition occurs is determined by the relative magnitude of the parameter γ. By incorporating (12) into (10), it follows that the resulting STVECM can display asymmetric behavior depending on the value of the transition variable, s_t. For example, one option, and one largely unexplored in the ERPT literature, is to set s_t equal to some observed measure of real economic activity such as the unemployment rate in an attempt to mimic the business cycle. Here we refer to a regression equation that uses a logistic transition function as a logistic smooth transition regression equation, or LSTR.

As specified in (10), it follows that each equation in the STVECM will share the same (identical) transition function. This is the approach most commonly applied in the literature; see, for example, Anderson and Vahid (1998), Rothman, van Dijk and Franses (2001) and Camacho (2004). From an empirical perspective such a specification may be overly restrictive. In other words, it is entirely possible that p_{it} will respond to s_t with a different speed than will p_{xt}. Of course it is even possible that the various equations in the system will have completely different transition functions, that is, some mix of logistic and exponential functions. In this spirit it is a straightforward matter to generalize (10) as follows:

$$\begin{aligned}\Delta \boldsymbol{y}_t &= (\boldsymbol{I} - \boldsymbol{\Gamma}_t)\left[\boldsymbol{\delta}_1 + \sum_{i=1}^{p-1} \boldsymbol{\Psi}_{i1} \Delta \boldsymbol{y}_{t-i} + \boldsymbol{\alpha}_1 \hat{\varepsilon}_{t-1}\right] \\ &+ \boldsymbol{\Gamma}_t \left[\boldsymbol{\delta}_2 + \sum_{i=1}^{p-1} \boldsymbol{\Psi}_{i2} \Delta \boldsymbol{y}_{t-i} + \boldsymbol{\alpha}_2 \hat{\varepsilon}_{t-1}\right] + \boldsymbol{\upsilon}_t, \qquad (13)\end{aligned}$$

where $\boldsymbol{I}$ is a 3×3 identity matrix and diag $(\boldsymbol{\Gamma}_t) = (G_1(s_{1t}), G_2(s_{2t}), G_3(s_{3t}))$, with off diagonal terms equalling zero. In this manner the STVECM in (10) may be generalized to allow for different transition functions (and transition variables) for each equation in the system. He, Teräsvirta and González (2008) considered a similar specification for a vector–autoregressive model, although they limited their analysis to the case where s_t simply equals the time index, t.

To our knowledge the STVECM framework has not been used to model regime dependent exchange rate pass–through effects. This is surprising given that the STVECM clearly nests many of the more common specifications used to examine nonlinear responses in the empirical literature on exchange rate pass–through.

4.2 A Testing Strategy: Single Equations

As is evident from both (10) and (13), the nonlinear features of the provisional STVECM model will depend on the selection of the transition function(s) as well as the transition variable(s). In practice there are typically a large number of options available during the model building phase. It is therefore desirable to have a testing strategy that reduces the number of nonlinear models that must ultimately be estimated and compared. To date there has been relatively little research on testing strategies for multivariate systems, with much of the focus being on testing in single equation models (Teräsvirta, 1994; Lundbergh, Terasvirta and van Dijk, 2003).

To gain insight into the testing problem, consider for the moment the case where (10) is reduced to a univariate smooth transition error correction model, that is, where $\boldsymbol{y}_t = \tilde{y}_t$ is a scalar. In this case we can re–write (10) as:

$$\Delta \tilde{y}_t = \boldsymbol{\varphi}_1' \tilde{\boldsymbol{x}}_t \left(1 - G(s_t; \boldsymbol{\theta})\right) + \boldsymbol{\varphi}_2' \tilde{\boldsymbol{x}}_t G(s_t; \boldsymbol{\theta}) + v_t, \tag{14}$$

where $\tilde{\boldsymbol{x}}_t = \left(1, \Delta \boldsymbol{y}_{t-1}', \ldots, \Delta \boldsymbol{y}_{t-p+1}', \hat{\varepsilon}_{t-1}\right)'$, a $(3{\times}p{+}1)$ vector, and where $\boldsymbol{\varphi}_1$ and $\boldsymbol{\varphi}_2$ are conformable parameter vectors. As well, assume that $G(.)$ is given by either (11) or (12). The problem, of course, is there are two ways to reduce (14) to a linear error correction model. On the one hand if $\boldsymbol{\varphi}_1 = \boldsymbol{\varphi}_2$, then the model becomes linear in parameters. Even so, it is not appropriate to simply test $\mathrm{H}_0 : \boldsymbol{\varphi}_1 = \boldsymbol{\varphi}_2$ given that in this case the γ and c parameters embedded in $G(.)$ are unidentified. Likewise, a standard test of $\mathrm{H}_0 : \gamma = 0$ is not appropriate given that in this case $\boldsymbol{\varphi}_1$ and $\boldsymbol{\varphi}_2$ are unidentified. The result in either case is the classical "Davies problem" outlined in a pair of papers by Davies (1977; 1987). The upshot is that tests of either null hypothesis will be associated with non–standard asymptotic distributions.

While various testing procedures have been proposed, a computationally convenient approach has been proposed by Luukkonen, Saikkonen and Teräsvirta (1988). Specifically, these authors advocate replacing the transition function $G(.)$ with a suitable Taylor series approximation, where the approximation is evaluated at $\gamma = 0$. If, for example, a third–order approximation is used, then a linear approximation to (14) is:

$$\Delta \tilde{y}_t = \boldsymbol{\psi}_1' \tilde{\boldsymbol{x}}_t + \boldsymbol{\psi}_2' \tilde{\boldsymbol{x}}_t s_t + \boldsymbol{\psi}_3' \tilde{\boldsymbol{x}}_t s_t^2 + \boldsymbol{\psi}_4' \tilde{\boldsymbol{x}}_t s_t^3 + \xi_t. \tag{15}$$

A test of linearity may now be conducted by simply testing $\mathrm{H}_0' : \boldsymbol{\psi}_2 = \boldsymbol{\psi}_3 = \boldsymbol{\psi}_4 = \mathbf{0}$ in (15). Note that while in general ξ_t contains both ε_t and approximation error, under the null hypothesis of linearity there is no approximation error. In this case $\varepsilon_t = \xi_t$, and standard Lagrange Multipler (LM) tests such including the F–test may be applied. That is, if RSS_1 denotes the error sum of squares from the restricted version of (15) and RSS_2 denotes the corresponding measure for the unrestricted model, then:

$$F_{LM} = \frac{(RSS_1 - RSS_2)/q}{RSS_2/(n-k)} \overset{approx}{\sim} F(q, T-p-1), \tag{16}$$

where $q = 3(p+1)$ are the number of restrictions implied by the null hypothesis H_0' and and k are the number of free parameters estimated in the unrestricted version of (15).

While the foregoing outlines a reasonable testing strategy for detecting nonlinearity, several issues remain. For example, it does not directly determine which transition function, that is, the exponential or the logistic, is most appropriate for a given application. Moreover, the nonlinearity test assumes that the transition variable, s_t, is known. While in some instances theory might dictate a likely candidate for transition variable, in many instances this choice, too, must be part of the overall testing framework. Regarding the first issue, Teräsvirta and Anderson (1992) and Teräsvirta (1994) describes a testing sequence that can be employed to identify the transition function. Specifically, assuming the linear model is rejected, the following conditional tests may be performed:

$$\mathrm{H}_{04}: \quad \boldsymbol{\psi}_4 = \mathbf{0}, \tag{17}$$

$$\mathrm{H}_{03}: \quad \boldsymbol{\psi}_3 = \mathbf{0} \mid \boldsymbol{\psi}_4 = \mathbf{0}, \tag{18}$$

$$\mathrm{H}_{02}: \quad \boldsymbol{\psi}_2 = \mathbf{0} \mid \boldsymbol{\psi}_3 = \boldsymbol{\psi}_4 = \mathbf{0}, \tag{19}$$

where again it is appropriate to use suitable F–versions of the tests implied by (17)–(19). The logic of the above testing sequence is that an exponential function is likely best approximated by a quadratic in s_t. Therefore, if (18) is rejected while (17) and (19) are not, the exponential function in (11) may be used. Alternatively, if (17) or (19) are rejected while (18) is not, than the logistic function in (12) may be tried.[8] Finally, there are few restrictions on candidates for the transition variable, s_t. Again, Teräsvirta (1994) suggests trying a slate of candidates and using the one associated with the strongest rejection of the linearity hypothesis, H_0'. Finally, once a candidate transition variable and transition function have been identified, provisional es-

[8] In the event that the testing sequence allows all hypotheses in (17)–(19) to be rejected, Teräsvirta (1994) suggests picking the transition function associated with the smallest p–value. For example, if a test of (18) yields the smallest p–value, an exponential transition function would be used.

timates of the smooth transition model in (14) can be obtained by employing nonlinear least squares (van Dijk, Teräsvirta and Franses, 2002). Furthermore, the diagnostic tests described by Eitrheim and Teräsvirta (1996) can be employed to examine model adequacy.

4.3 A Testing Strategy: Multivariate Systems

As noted previously, there is a paucity of studies that have explored nonlinearity testing in a multivariate setting, especially when a system such as (13) is examined with equation–specific transition functions. Even so, Rothman, van Dijk and Franses (2001), Camacho (2004), and Péguin-Feissolle, Strikholm and (2008) are notable exceptions, with each of these studies advancing a framework for testing nonlinearities in a multi–equation model. In principle doing so is straightforward: the multivariate counterpart to (15) may be specified as:

$$\Delta \boldsymbol{y}_t = \boldsymbol{F}_1 \boldsymbol{X}_t + \boldsymbol{F}_2 \boldsymbol{X}_t \boldsymbol{s}_{1t} + \boldsymbol{F}_3 \boldsymbol{X}_t \boldsymbol{s}_{2t} + \boldsymbol{F}_4 \boldsymbol{X}_t \boldsymbol{s}_{3t} + \boldsymbol{\xi}_t,\ \boldsymbol{\xi}_t \sim N\left(\mathbf{0}, \boldsymbol{\Sigma}\right), \tag{20}$$

where in this case X_t is a $3 \times (p+1)$ matrix defined as $\boldsymbol{X}_t = \boldsymbol{\iota} \tilde{\boldsymbol{x}}_t'$, and where $\boldsymbol{\iota}$ is a (3×1) unit vector. As well, $\boldsymbol{s}_{it} = \left(s^i_{1t}, s^i_{2t}, s^i_{3t}\right)'$, $i = 1, 2, 3$, $\boldsymbol{F}_i$, $i = 1, \ldots, 4$, are conformable parameter matrices, and where $\boldsymbol{\Sigma}$ is a symmetric, positive–definite error covariance matrix. The system nonlinearity test then involves a test of the hypothesis $\mathrm{H}_0'' : \boldsymbol{F}_2 = \boldsymbol{F}_3 = \boldsymbol{F}_4 = 0$, which will involve $q = 3\left[3 \times (p+1)\right]$ linear restrictions on the parameters of (20).

Following Bewley (1986), an F–version of the LM test of H_0'' in the multi–equation system is:

$$F_{LMS} = \frac{T}{q}\left(m - \mathrm{tr}\left(\hat{\boldsymbol{\Omega}}_1 \hat{\boldsymbol{\Omega}}_0^{-1}\right)\right) \overset{approx}{\sim} F\left(q, T\right), \tag{21}$$

where $\boldsymbol{\Omega}_0 = T\hat{\boldsymbol{\Sigma}}_0$ with $\hat{\boldsymbol{\Sigma}}_0$ being the estimated residual covariance matrix for the model under the null, $\boldsymbol{\Omega}_1 = T\hat{\boldsymbol{\Sigma}}_1$ similarly defined for the model under the alternative, and m is the number of equations in the system (here $m = 3$). While a value for F_{LMS} that exceeds the critical value from the $F\left(q, T\right)$ distribution is a clear indication of nonlinearity in the system, it says nothing about which equation(s) are appropriately nonlinear, nor does it suggest which transition function or set of transition variables are most applicable.

In principle a multivariate version of the testing sequence in (17)–(19) could be also performed.

While as such a richer, more detailed sequence of tests could be developed, the fact is the number of combinations of candidate transition variables and transition functions involved could quickly become overwhelming. We therefore propose a simple yet practical strategy for identifying the appropriate form of the STVECM in (13). Specifically, we propose using the single–equation testing framework outlined in the previous section for specifying the structure of each equation in the system. Furthermore, once a set of candidate transition variables has been identified, the test in (21) may be employed to evaluate system–wide nonlinearity.

5 Empirical Results

5.1 Nonlinearity Testing Results

The testing and estimation methods described above are used to examine nonlinearity in exchange rate pass–through for U.S. and Canadian OSB prices. The approach first necessitates estimating a best–fitting linear error correction model for each equation. The explanatory variables used are lags of (first differences) of representative (logarithmic) OSB import and export prices and the first difference of the log of the U.S. dollar–Canadian dollar exchange rate. A systems version of Akaike's information criterion (AIC) is used to determine appropriate lag lengths.[9] The AIC indicated that up to four lags of the $\Delta \boldsymbol{y}_t$ vector are needed in each equation. Even so, two additional lags were called for to render the residuals of the foreign exchange equation white noise. Additional testing confirmed that exchange rates respond only to their own lags, and are therefore exogenous to OSB prices. As well, preliminary tests suggested that lagged changes in exchange rates are insignificant in the OSB price equations.[10]

The results of nonlinearity tests applied to the U.S. and Canadian OSB price equations are reported in Table 2. Candidates for the transition variables include up to six lags of the lagged residual from the estimated cointegrating equation (i.e., $\hat{\varepsilon}_{t-j},\ j = 1, \ldots, 6$) and a 52–week moving

[9] Specifically, we use AIC $= ln\left(\det(\hat{\Sigma})\right) + 2N/T$, where N denotes the number of estimated parameters in the model.

[10] Of course this result does not preclude the possibility of ERPT into OSB prices, as the lagged cointegrating residuals, which incorporate the lagged exchange rate, remain in the specifications.

average of the unemployment rate, that is,

$$\widetilde{une}_t \;=\; \frac{1}{52}\sum_{i=1}^{52} une_{t-i}. \tag{22}$$

The fifty–two week average smooths out short–term and seasonal fluctuations in the weekly unemployment rate, and therefore should over time send a reasonable signal of general economic conditions. The test results show that for both equations, linearity is most convincingly rejected for the $\widetilde{une}_t$ variable. Moreover, the results of applying the testing sequence in (17)–(19) suggest that the transition function is likely a logistic as specified in (12). The implication, then, is that ERPT into OSB prices is likely asymmetric, and moreover that this asymmetry occurs in conjunction with a general indicator of the business cycle. This preliminary result is, moreover, consistent with recent work by Chew, Ouliaris and Tan (2011).

At this stage several additional issues must be considered. First is the question of what transition variable is most likely associated with nonlinearity in the exchange rate equation, which contains an intercept and six lags of the log difference of exchange rates. To this end, the nonlinearity tests were repeated for the exchange rate equation; the results are reported in the left–hand panel of Table 3. Of the transition variables considered, results in Table 3 indicate the presence of substantial nonlinearities in the exchange rate equation, with $s_t = \Delta e_{t-1}$ being associated with the strongest rejection of linearity. And for this variable the testing sequence suggests that an LSTR might be the most appropriate specification, although the rejection of H_{03} is also quite strong, indicating that an ESTR specification could also be acceptable.

The preliminary evidence reported above suggests that a 52–week moving average of unemployment is a reasonable transition variable in both OSB price equations. Therefore, it may be desirable to incorporate a fourth equation into the system to explain weekly unemployment rates. Moreover, prior work–see, for example, van Dijk, Teräsvirta and Franses (2002) and Deschamps (2008)–has found substantial evidence in favor of LSTR models for monthly U.S. unemployment rates. Even so, to our knowledge prior studies have not focused on modeling unemployment rates (based on unemployment claims) on a weekly basis. The base linear model used here is of the form:

$$\Delta\widetilde{y}_t = \lambda_0 + \sum_{i=1}^{3}\left(\eta_i \sin\left(2\pi t/f_i\right) + \kappa_i \cos\left(2\pi t/f_i\right)\right) + \sum_{i=1}^{p-1}\lambda_i \Delta\widetilde{y}_{t-i} + \rho y_{t-1} + \upsilon_t, \tag{23}$$

where $\widetilde{y}_t = une_t$ and $f_1 = 13$, $f_2 = 26$, and $f_3 = 52$. The sine–cosine terms are incorporated to account for the seasonal nature of unemployment claims. As well, we follow Skalin and Teräsvirta (2002) by including a lagged level term for the unemployment variable, which in turn implies that unemployment follows a "natural rate" (i.e., is mean reverting) as opposed to a "hysteresis" hypothesis.[11] Of course once nonlinearities are considered, it is possible that unemployment rates could even display locally explosive behavior.

The linear model in (23) was fitted to the data. The (univariate) AIC indicated that up to eleven lags of Δune_t are needed to eliminate residual serial correlation. Results of applying linearity tests for the unemployment rate equation are recorded in the right–hand panel of Table 3. Consistent with prior studies, as well as with the asymmetries that may be detected by simply the data plot in Figure 3, there is overwhelming evidence of nonlinearity in the unemployment data. Results in Table 3 suggest that linearity is rejected most convincingly for$(une_{t-1} - une_{t-4})$. Of interest is that the seasonal difference $(une_{t-1} - une_{t-53})$ and the 52–week moving average $\widetilde{une}_t$, while indicating the presence of nonlinearities, are not the strongest candidates for a transition variable in the unemployment equation.[12] In all instances the testing sequence overwhelmingly indicates that an LSTR model is called for, a result that is, moreover, also consistent with prior research (van Dijk, Teräsvirta and Franses, 2002).

5.2 Smooth Transition Model Results

The foregoing suggests there is evidence of nonlinearity in each equation in the system, which among other things suggests that ERPT into OSB prices may have a regime–dependent effect. As discussed in Section 4, as part of the STVECM model building process we first estimate suitable univariate smooth transition models for each equation.

The results of the univariate analysis are summarized in Table 4–there we report model fit diagnostic measures for each of the estimated linear and nonlinear models for each variable in the system. To begin, preliminary estimations revealed that an LSTR specification for the exchange

[11] Of course the results reported in Table 1 suggest that une_t behaves in a manner consistent with a unit root process (i.e., hysteresis). Even so, Skalin and Teräsvirta (2002) report that it is often difficult reject the null of a unit root even when the underlying data were generated in a manner consistent with mean–reverting behavior and strong asymmetries.

[12] Alternatively, when using monthly U.S. unemployment data van Dijk, Teräsvirta and Franses (2002) and Deschamps (2008) find that a lagged seasonal difference works quite well as a transition variable.

rate equation (that uses $s_t = \Delta y_{t-1}$ as a transition variable) ended up fitting only a small handful of outliers. Given the results in Table 3, we also fitted an ESTR to the exchange rate series, which yielded more satisfactory results. Turning to an assessment of the univariate models, results in Table 4 show that in every case the nonlinear model represents an improvement in fit relative to its linear counterpart, with the nonlinear unemployment equation yielding the biggest increase in fit relative to its linear counterpart and the exchange rate equation the smallest. In addition, there is little evidence of remaining autocorrelation in each model's residuals up to a twelve–week lag (the smooth transition model for unemployment at lags six and twelve being an exception). Results in Table 4 also indicate that the residuals for each estimated model are highly leptokurtic (i.e., they are associated with "fat tails"), which is not surprising given the relatively high frequency of the data (weekly). There is also evidence of ARCH errors in each case, a result that, moreover, might be anticipated given the weekly frequency of the data.

As a final check of the nonlinear specifications, the system nonlinearity test, as outlined in (21), was applied to the four–equation system. In conducting the test the system in (20) was estimated where the transition variables identified for the univariate models in Table 4 are used. The resulting test statistic, 2.690, is extreme in the corresponding $F_{(132,607)}$ distribution. Taken together, this result and those recorded in Table 4 suggest that nonlinearity is an important feature of these data.

The final step in constructing a model for assessing regime–dependent ERPT into North American OSB prices is to estimate the STVECM. The transition functions and transition variables used in specifying the univariate models are maintained; the parameter estimates obtained for the univariate models are used as starting values. The system estimation results, along with several summary measures of model fit, are reported in Table 5. Plots of the corresponding estimated transition functions for each equation, both over time and with respect to each implied transition variable, are reported in Figure 4. Additional tests revealed that covariance terms amongst the price variables and exchange rates and unemployment were not significantly different from zero, as is the covariance term between the exchange rate and unemployment. These restrictions are incorporated in the estimates recorded for the STVECM reported here.

As indicated in Table 5, the STVECM provides a substantial improvement in fit relative to the linear VECM; for example, the ratio of the determinant for the STVECM's covariance matrix relative to its linear counterpart is 0.603. As well, the system AIC also indicates an improvement in fit for the STVECM relative to the linear VECM. Regarding the implied nonlinearities, the plots in Figure 4 show that, with the exception of the transition function for the OSB price in Eastern Canada, the estimated transition functions imply a smooth response to changes in the respective transition variables. The plots in Figure 4 also suggest that the transition functions for the OSB price equations, when plotted over time, do a reasonable job of tracking recent business cycle behavior. Finally, the parameter estimates reported in Table 5 suggest that, for each estimated equation, the estimated parameters change substantially with respect to the implied transition functions, including the speed–of–adjustment parameters associated with the lagged error correction terms in the OSB price equations. Furthermore, the STVECM apparently does a reasonable job of generating results for prices, the exchange rate, and the unemployment rate that are consistent with observed behavior. Along with the observed data, Figures 2 and 3 show the realizations of a single Monte Carlo simulation of the model from the end of the sample period (August, 2010) through the middle of 2014. In each case the simulated data seemingly depicts various features of the observed data, including asymmetries. Taken together, the results for the estimated STVECM suggest there is scope for ERPT into OSB prices to vary with the weekly U.S. unemployment rate and that, moreover, unemployment itself is also a highly nonlinear process.

5.3 Generalized Impulse Response Functions

To assess the effects of ERPT into OSB prices, it is useful to generate generalized impulse response functions (GIRFs). Specifically, Koop, Pesaran and Potter (1996) define a set of procedures that may be applied to compute GIRFs for multivariate nonlinear models. A (multivariate) GIRF is defined by:

$$G_{\Delta \boldsymbol{y}}\left(n, \boldsymbol{\delta}, \boldsymbol{\omega}_{t-1}\right) = E\left(\Delta \boldsymbol{y}_{t+n} \left| \boldsymbol{v}_t = \boldsymbol{\delta}, \Omega_{t-1} = \omega_{t-1}\right.\right) - E\left(\Delta \boldsymbol{y}_{t+n} \left| \boldsymbol{v}_t = \mathbf{0}, \Omega_{t-1} = \boldsymbol{\omega}_{t-1}\right.\right), \qquad (24)$$

where n denotes the forecast horizon, $\boldsymbol{\delta}$ is a vector of shocks, $\Omega_{t-1} = \boldsymbol{\omega}_{t-1}$ denotes information available through period $t-1$ (i.e., the history), and E is an expectation operator. To determine the initial conditions, we randomly draw (with replacement) 50 histories (i.e., ω_{t-1}'s) from the set of 607 available histories. As is common in the ERPT literature, we then consider unit shocks to the exchange rate equation (Cashin, Liang and McDermott, 2000) and, as well, to the unemployment rate. To evaluate the expectations in (24), we use 600 Monte Carlo draws from a multivariate random normal distribution with a variance–covariance matrix equal to that of the estimated STVECM. Impulse responses for the levels of the variables in the system are computed by summing those obtained for the first differences, that is, by constructing:

$$G_{\boldsymbol{y}}\left(n, \boldsymbol{\delta}, \boldsymbol{\omega}_{t-1}\right) = \sum_{i=1}^{n} G_{\Delta \boldsymbol{y}}\left(n, \boldsymbol{\delta}, \boldsymbol{\omega}_{t-1}\right). \tag{25}$$

Finally, it is also possible to construct regime–dependent GIRFs where, for example, shocks can be initiated only when $G_1(s_{1t}) \geq 0.5$ or $G_1(s_{1t}) < 0.5$.[13] In this manner it is possible to examine the extent to which ERPT into OSB prices varies with the unemployment rate.

Unconditional GIRFs for a one–time unit shock (both positive and negative) to the U.S. dollar–Canadian dollar exchange rate, taken over a 156–week horizon, are reported in Figure 5. As illustrated there, pass–through of such a shock into the U.S. OSB price is never complete, reaching at most 75–percent. Moreover, the effects are initially quite small–they do not reach even 50–percent during the first year following the shock. As well, the GIRFs appear to be symmetric with respect to positive versus negative exchange rate shocks. This result is reasonable given that: (1) unemployment is not impacted by nominal exchange rate movements (and therefore there is no systematic "regime change" for the OSB price equations); and (2) that nonlinearity in the exchange rate equation is associated with an ESTR, which is (nearly) symmetric around zero.

A different picture emerges, however, when conditional GIRFs are computed for an exchange rate shock; see Figure 6. As the figure shows, when the 52–week moving average of unemployment (i.e., s_{1t}) is greater than 2.91–percent, that is, when $G_1(s_{1t}) \geq 0.5$, ERPT associated with a positive one–unit shock reaches unity (i.e., is complete) after only twelve weeks. Indeed, as depicted in Figure 6, this stabilizes at a value far in excess of unity–near three, in fact–after approximately two years have elapsed. Conversely, the GIRFs conditional on the moving average of unemployment being less than 2.91–percent (i.e., $G_1(s_{1t}) < 0.5$) illustrate that pass–through is, again, slow to respond and, moreover, relatively incomplete, even after three years have elapsed; the long–run response to a positive unit shock in this case is about 0.44–percent. These results firmly establish that ERPT into prices for a primary home construction material, that is, oriented strand board, is highly regime

[13] Given the estimate for the centrality parameter, c_1, reported in Table 5, the conditional GIRFs in this case are consistent with the 52–week moving average of unemployment rates, $\widetilde{une}_t$, being above or below 2.908.

dependent and that, moreover, the regimes themselves are a function of the overall performance of the general economy.

Because of the nature of the model it is also possible to obtain GIRFs associated with an unemployment shock, in this case with respect to a one standard deviation shock to the unemployment rate. The resulting unconditional GIRFs are reported in Figure 7. They show, for example, that a positive shock to unemployment apparently causes unemployment rates themselves to continue to rise throughout the three–year horizon. As well, the impact on OSB prices is initially positive but after approximately 100 weeks the effects become negative. Furthermore, the GIRFs in Figure 7 that the estimated STVECM is apparently not dynamically stable with respect to unemployment shocks, as the GIRFs for prices and unemployment do not stabilize at a new level. This is not the case, however, as revealed in Figure 8. There we see the GIRF for unemployment (in response to an unemployment rate shock) extended over a six–year horizon. What is revealed there is that after approximately six years have elapsed that the GIRFs for unemployment effectively return to zero. In short, the model depicts something akin to a six–year peak–to–peak business cycle (based on unemployment), a result that is, moreover, in keeping with the general conclusion that post–war business cycles in the United States have lasted, on average, for approximately five–six years (see, .e.g., Watson, 1994).

As before, it is possible to obtain conditional GIRFs for unemployment shocks, in this case when $G_4(s_{4t}) \geq$ (respectively $<$) 0.5. The results for these conditional GIRFs are reported in Figure 9. Among other things the plots in Figure 9 bring into focus the asymmetries associated with unemployment. To begin, the effects on the U.S. price for OSB associated with a positive shock are, as before, initially positive, although the peak occurs much more quickly when $s_{t4} = une_{t-1} - unet - 4 \geq 0.149$, that is, when unemployment rates are trending higher (25 weeks versus approximately 45 weeks). Even so, the effects resulting from a positive shock apparently converge after approximately 100 weeks. More interesting, however, is that negative shocks yield GIRFs that are always larger in absolute terms when $G_4 s_{4t} \geq 0.5$ as opposed to when $G_4 s_{4t} \geq 0.5$. Apparently a decrease in unemployment has a larger effect on OSB prices when unemployment rates are, relatively speaking, already high than when the converse is true. Similar results occur for the conditional GIRFs associated with the Canadian OSB price associated with an unemployment shock.

6 Summary and Conclusions

In this study we have examined exchange rate pass–through into oriented strand board, an important construction material produced and traded throughout much of North America. Indeed, Canada and the United States are leading producers of OSB, but historically Canada has exported more than 75–percent of its total OSB production to the United States. In the U.S. OSB is produced primarily in the Southeastern region of the country, although in recent decades this region has also experienced the most rapid growth (and, since 2007, the most rapid declines) in new home construction. To investigate ERPT into OSB prices, we obtained weekly mill–gate prices from *Random Lengths* for the 1998–2010 period. Specifically, the prices correspond to mill prices for OSB in Eastern Canada (prices for mills in Ontario and Quebec) and the Southeast U.S. (prices for mills in Georgia, Alabama, Mississippi, South Carolina, and Tennessee). Furthermore, the Canadian prices are recorded in U.S. dollars, that is, local currency pricing is employed.

Recent work by Goodwin, Holt and Prestemon (2011) found evidence of nonlinearity in the LOP relationship between these prices, but they did not consider ERPT effects. Moreover, recent research has examined nonlinear and asymmetric ERPT into import prices by assuming that deviations from the underlying long–run equilibrium relationship will have a differential impact on estimated pass–through responses depending on the overall dagnitude of the deviations (see, e.g., Al-Abri and Goodwin, 2009; Larue, Gervais and Rancourt, 2010). More recently, several authors have investigated asymmetric effects of ERPT into prices as a function of overall macroeconomic activity (Nogueira, Jr. and León-Ledesma, 2011; Chew, Ouliaris and Tan, 2011), albeit for aggregate price indices and not for specific industries or commodity prices.

Building on prior work in this general area, we examine the asymmetric effects of long–term swings in weekly unemployment claims on ERPT into prices for OSB. We do so by proposing a feasible strategy for building and estimating a smooth transition vector error correction model wherein each equation is allowed to have its own built–in asymmetries (i.e., transition function and transition variables). Specifically, we estimate a four–equation STVECM where asymmetries in the the OSB price equations are modeled by using logistic transition functions where, moreover, the transition variables are in both cases a 52–week moving average of the unemployment rate. Nonlinearities in the nominal U.S. dollar–Canadian dollar exchange rate are modeled by using an exponential transition function. And finally, in a manner consistent with prior work on modeling asymmetries in unemployment rates (see, e.g., Skalin and Teräsvirta, 2002), we model asymmetries in weekly unemployment rates by using a logistic smooth transition model.

An immediate implication of the estimated STVECM is as follows: not only is there the potential for direct asymmetric (nonlinear) ERPT into OSB prices, but also the potential for indirect effects due to the regime–dependent behavior identified separately for the exchange rate and unemployment equations. To our knowledge no prior study has allowed for such a rich specification of nonlinearities when examining ERPT. To assess the nature of nonlinearities in ERPT, we employ the generalized impulse response function framework of Koop, Pesaran and Potter (1996). Similar to prior work on this general topic, we find incomplete pass–through into OSB prices in both the U.S. and Canada. Moreover, for OSB prices in the Southeast U.S., pass–through effects are very small in the short run and only reach a long–run steady state of approximately 0.75 (for a positive shock) after more than two years have elapsed. As well, these estimated effects depend in a striking way on overall macroeconomic conditions. Specifically, conditional GIRFs, that is, GIRFs obtained for when unemployment rates are high versus low, indicate that ERPT effects during the recent economic downturn were far greater than unity. These results are, moreover, in keeping with prior work by Nogueira, Jr. and León-Ledesma (2011) and Chew, Ouliaris and Tan (2011), who also find that pass–through effects increase substantially during economic downturns. Finally, because of the way the STVECM is specified, we can also examine GIRFs associated with unemployment shocks, that is, unemployment pass–through effects. We find these effects are generally smaller than for exchange rate shocks, although they vary considerably over a six–seven year period, which in turn is roughly consistent with the observed span for post–war business cycle activity.

While this paper represents an important contribution to the ERPT literature, and especially so for timber products, more work remains. Specifically, it would be useful to examine the potential asymmetric responses to other measures of macroeconomic activity. As well, it would be desirable to repeat the analysis for price relationships involving other products and trade to–from other countries and regions. Even so, we believe the work reported here provides a good starting point for subsequent studies on these and related topics.

References

Al-Abri, Almukhtar S. and Barry K. Goodwin, "Re–examining the Exchange Rate Pass–through into Import Prices using Non–linear Estimation Techniques: Threshold Cointegration," *International Review of Economics & Finance*, January 2009, *18* (1), 142–161.

Anderson, Heather M. and Farshid Vahid, "Testing Multiple Equation Systems for Common Nonlinear Components," *Journal of Econometrics*, May 1998, *84* (1), 1–36.

Bailliu, Jeannine and Eiji Fujii, "Exchange Rate Pass–Through and the Inflation Environment in Industrialized Countries: An Empirical Investigation," Working Papers 04-21, Bank of Canada 2004.

Balke, Nathan S. and Thomas B. Fomby, "Threshold Cointegration," *International Economic Review*, August 1997, *38* (3), 627–645.

Bewley, Ron, *Allocation Models*, Cambridge, Massachusetts: Ballinger Publishing Company, 1986.

Bolkesjø, Torjus F. and Joseph Buongiorno, "Short- and Long-Run Exchange Rate Effects on Forest Product Trade: Evidence from Panel Data," *Journal of Forest Economics*, 2006, *11* (4), 205–221.

Bussiere, Matthieu, "Exchange Rate Pass–Through to Trade Prices – The Role of Non–Linearities and Asymmetries," Working Paper Series 822, European Central Bank October 2007.

Camacho, Maximo, "Vector Smooth Transition Regression Models for US GDP and the Composite Index of Leading Indicators," *Journal of Forecasting*, 2004, *23* (3), 173–196.

Campa, José Manuel and Linda S. Goldberg, "Exchange Rate Pass–Through into Import Prices," *The Review of Economics and Statistics*, December 2005, *87* (4), 679–690.

Campa, José Manuel, Linda S. Goldberg, and Jose M. Gonzalez-Minguez, "Exchange Rate Pass–Through to Import Prices in the Euro Area," Staff Reports 219, Federal Reserve Bank of New York 2005.

Cashin, Paul, Hong Liang, and C. John McDermott, "How Persistent Are Shocks to World Commodity Prices?," *IMF Staff Papers*, 2000, *47* (2), 177–217.

Chew, Joey, Ouliaris, and Siang Meng Tan, "Exchange Rate Pass–Through over the Business Cycle in Singapore," IMF Working Papers 11/141, International Monetary Fund June 2011. http://ideas.repec.org/p/imf/imfwpa/11-141.html.

Davies, R. B., "Hypothesis Testing When a Nuisance Parameter is Present Only Under the Alternative," *Biometrika*, 1977, *64* (2), pp. 247–254.

Davies, Robert B., "Hypothesis Testing when a Nuisance Parameter is Present Only Under the Alternatives," *Biometrika*, 1987, *74* (1), pp. 33–43.

Demetrescu, Matei, "On the Dickey–Fuller Test with White Standard Errors," *Statistical Papers*, January 2010, *51* (1), 11–25.

Deschamps, Philippe J., "Comparing Smooth Transition and Markov Switching Autoregressive Models of US Unemployment," *Journal of Applied Econometrics*, 2008, *23* (4), 435–462.

Dickey, David A. and Wayne A. Fuller, "Distribution of the Estimators for Autoregressive Time Series With a Unit Root," *Journal of the American Statistical Association*, 1979, *74* (366), pp. 427–431.

Eitrheim, Øyvind and Timo Teräsvirta, "Testing the Adequacy of Smooth Transition Autoregressive Models," *Journal of Econometrics*, September 1996, *74* (1), 59–75.

Gagnon, Joseph E. and Jane Ihrig, "Monetary Policy and Exchange Rate Pass–Through," *International Journal of Finance & Economics*, 2004, *9* (4), 315–338.

Goldberg, Pinelopi Koujianou and Michael M. Knetter, "Goods Prices and Exchange Rates: What Have We Learned?," *Journal of Economic Literature*, September 1997, *35* (3), 1243–1272.

Goodwin, Barry K., Matthew T. Holt, and Jeffrey P. Prestemon, "North American Oriented Strand Board Markets, Arbitrage Activity, and Market Price Dynamics: A Smooth Transition Approach," *American Journal of Agricultural Economics*, 2011, *93* (4), 993–1014.

He, Changli, Timo Teräsvirta, and Andreas González, "Testing Parameter Constancy in Stationary Vector Autoregressive Models Against Continuous Change," *Econometric Reviews*, 2008, *28* (1–3), 225–245.

Jabara, Cathy L. and Nancy E. Schwartz, "Flexible Exchange Rates and Commodity Price Changes: The Case of Japan," *American Journal of Agricultural Economics*, 1987, *69* (3), 580–590.

Karoro, Tapiwa D., Meshach J. Aziakpono, and Nicolette Cattaneo, "Exchange Rate Pass–Through to Import Prices in South Africa: Is there Asymmetry?," *South African Journal of Economics*, 2009, *77* (3), 380–398.

Knetter, Michael M., "Price Discrimination by U.S. and German Exporters," *The American Economic Review*, 1989, *79* (1), 198–210.

Koop, Gary, M. Hashem Pesaran, and Simon M. Potter, "Impulse Response Analysis in Nonlinear Multivariate Models," *Journal of Econometrics*, September 1996, *74* (1), 119–147.

Larue, Bruno, Jean-Philippe Gervais, and Yannick Rancourt, "Exchange Rate Pass–Through, Menu Costs and Threshold Cointegration," *Empirical Economics*, 2010, *38*, 171–192. 10.1007/s00181-009-0261-2.

Leamer, Edward E., "Housing IS the Business Cycle," NBER Working Papers 13428, National Bureau of Economic Research, Inc September 2007. http://ideas.repec.org/p/nbr/nberwo/13428.html.

Lundbergh, Stefan, Timo Terasvirta, and Dick van Dijk, "Time–Varying Smooth Transition Autoregressive Models," *Journal of Business & Economic Statistics*, January 2003, *21* (1), 104–21.

Luukkonen, Ritva, Pentti Saikkonen, and Timo Teräsvirta, "Testing Linearity Against Smooth Transition Autoregressive Models," *Biometrika*, 1988, *75* (3), 491–499.

MacKinnon, James G., "Critical Values for Cointegration Tests," Working Papers 1227, Queen's University, Department of Economics January 2010.

Magee, Lonnie, "R2 Measures Based on Wald and Likelihood Ratio Joint Significance Tests," *The American Statistician*, 1990, *44* (3), 250–253.

Mann, Catherine L., "Prices, Profit Margins, and Exchange Rates," *Federal Reserve Bulletin*, 1986, (Jun), 366–379.

Marazzi, Mario and Nathan Sheets, "Declining Exchange Rate Pass–through to U.S. Import Prices: The Potential Role of Global Factors," *Journal of International Money and Finance*, 2007, *26* (6), 924–947.

Ng, Serena and Pierre Perron, "Unit Root Tests in ARMA Models with Data-Dependent Methods for the Selection of the Truncation Lag," *Journal of the American Statistical Association*, 1995, *90* (429), 268–281.

Nogueira, Jr., Reginaldo P. and Miguel A. León-Ledesma, "Does Exchange Rate Pass–Through Respond to Measures of Macroeconomic Instability?," *Journal of Applied Economics*, 2011, *XIV*, 167–180.

Parsley, David C, "Exchange–Rate Pass–through with Intertemporal Linkages: Evidence at the Commodity Level," *Review of International Economics*, October 1995, *3* (3), 330–341.

Péguin-Feissolle, Anne, Birgit Strikholm, and Timo Teräsvirta, "Testing the Granger Non-causality Hypothesis in Stationary Nonlinear Models of Unknown Functional Form," CREATES Research Papers 2008–19, School of Economics and Management, University of Aarhus April 2008.

Phillips, Peter C. B. and Pierre Perron, "Testing for a Unit Root in Time Series Regression," *Biometrika*, 1988, *75* (2), 335–346.

Pollard, Patricia S. and Cletus C. Coughlin, "Size Matters: Asymmetric Exchange Rate Pass–Through at the Industry Level," Working Papers 2003-029, Federal Reserve Bank of St. Louis 2004.

Rothman, Philip, Dick van Dijk, and Philip Hans Franses, "Multivariate Star Analysis Of Money Output Relationship," *Macroeconomic Dynamics*, September 2001, *5* (04), 506–532.

Skalin, Joakim and Timo Teräsvirta, "Modeling Asymmetries And Moving Equilibria In Unemployment Rates," *Macroeconomic Dynamics*, April 2002, *6* (02), 202–241.

Stock, James H. and Mark M. Watson, "How Did Leading Indicator Forecasts Perform During the 2001 Recession?," *Economic Quarterly*, 2003, (Sum), 71–90.

Teräsvirta, Timo, "Specification, Estimation, and Evaluation of Smooth Transition Autoregressive Models," *Journal of the American Statistical Association*, 1994, *89* (425), 208–218.

Teräsvirta, Timo and Heather M. Anderson, "Characterizing Nonlinearities in Business Cycles Using Smooth Transition Autoregressive Models," *Journal of Applied Econometrics*, 1992, *7*, S119–S136.

Uusivuori, Jussi and Joseph Buongiorno, "Pass–Through of Exchange Rates on Prices of Forest Product Exports from the United States to Europe and Japan," *Forest Science*, 1991, *37* (3), 931–948.

van Dijk, Dick, Timo Teräsvirta, and Philip Hans Franses, "Smooth Transition Autoregressive Models – A Survey of Recent Developments," *Econometric Reviews*, 2002, *21* (1), 1–47.

Watson, Mark W., "Business-Cycle Durations and Postwar Stabilization of the U.S. Economy," *The American Economic Review*, 1994, *84* (1), 24–46.

Webber, Anthony G., "Newton's Gravity Law and Import Prices in the Asia Pacific," *Japan and the World Economy*, January 2000, *12* (1), 71–87.

Wickremasinghe, Guneratne Banda and Param Silvapulle, "Exchange Rate Pass–Through to Manufactured Import Prices: The Case of Japan," International Trade 0406006, EconWPA June 2004.

Table 1: Unit Root and Cointegration Test Results for OSB Pass–Through Data.

Variable	ADF	PP
Unit Root:		
p_i	-2.048	-2.663
p_x	-2.967	-2.882
e	-0.992	-0.833
une	-1.891	-1.887
Critical Values:		
1-percent	-3.444	-3.444
5-percent	-2.867	-2.867
Cointegration:		
	-6.902	-6.703
Critical Values:		
1-percent	-4.318	-4.318
5-percent	-3.755	-3.755
10-percent	-3.462	-3.462

Note: p_i denotes the import price (Southeast U.S.); p_x the export price (Eastern Canada); e the nominal U.S. dollar/Canadian dollar exchange rate; and une the unemployment rate. The column headed ADF reports heteroskedasticity robust Augmented Dickey–Fuller test statistics. The column headed PP denotes Phillips–Perron unit root test statistics. Results labeled cointegration are for a unit root test of the residuals of an Engle–Granger cointegrating regression of the import price on the export price and the exchange rate. All critical values were obtained from MacKinnon (2010).

Table 2: Single–Equation Nonlinearity Test Results for Southeast U.S. and North Eastern Canada.

	Import Price–Southeast U.S.				Export Price–Eastern Canada			
s_t	H_0'	H_{04}	H_{03}	H_{02}	H_0'	H_{04}	H_{03}	H_{02}
$\hat{\varepsilon}_{t-1}$	0.102	0.676	0.240	0.028	0.024	0.768	0.065	0.010
$\hat{\varepsilon}_{t-2}$	0.213	0.504	0.792	0.025	0.220	0.170	0.514	0.262
$\hat{\varepsilon}_{t-3}$	0.437	0.789	0.400	0.177	0.511	0.810	0.308	0.320
$\hat{\varepsilon}_{t-4}$	0.198	0.494	0.466	0.068	0.070	0.559	0.127	0.048
$\hat{\varepsilon}_{t-5}$	0.356	0.851	0.158	0.254	0.039	0.902	0.010	0.072
$\hat{\varepsilon}_{t-6}$	0.294	0.188	0.165	0.841	0.056	0.470	0.021	0.247
$\widetilde{une}_t$	0.004	0.157	0.926	6.82×10^{-5}	0.006	0.015	0.656	0.010

Note: The column headed s_t defines the candidate transition variable. Entries are are approximate p–values for the LM tests of nonlinearity (H_0'), and for the sequence of tests defined in (17)–(19) for determining whether the transition function is likely an exponential or a logistic.

Table 3: Single–Equation Nonlinearity Test Results for Exchange Rate and Unemployment Rate.

	Exchange Rate					Unemployment Rate			
s_t	H_0'	H_{04}	H_{03}	H_{02}	s_t	H_0'	H_{04}	H_{03}	H_{02}
Δe_{t-1}	1.71×10^{-15}	1.04×10^{-3}	4.14×10^{-6}	6.36×10^{-10}	Δune_{t-1}	2.78×10^{-24}	1.60×10^{-3}	2.90×10^{-4}	4.83×10^{-23}
$(e_{t-1}-e_{t-3})$	9.64×10^{-11}	1.70×10^{-3}	1.54×10^{-3}	1.16×10^{-7}	$(une_{t-1}-une_{t-3})$	3.67×10^{-22}	2.80×10^{-3}	0.029	6.63×10^{-24}
$(e_{t-1}-e_{t-4})$	2.86×10^{-12}	0.013	0.026	7.52×10^{-12}	$(une_{t-1}-une_{t-4})$	3.10×10^{-25}	0.013	9.70×10^{-5}	5.25×10^{-25}
$(e_{t-1}-e_{t-5})$	4.87×10^{-8}	0.095	7.33×10^{-3}	2.05×10^{-7}	$(une_{t-1}-une_{t-5})$	1.75×10^{-20}	0.041	1.57×10^{-6}	6.70×10^{-18}
$(e_{t-1}-e_{t-6})$	1.14×10^{-10}	0.132	2.57×10^{-9}	5.25×10^{-4}	$(une_{t-1}-une_{t-6})$	1.46×10^{-24}	8.72×10^{-5}	0.022	2.27×10^{-24}
$(e_{t-1}-e_{t-7})$	7.04×10^{-12}	0.041	5.25×10^{-7}	6.65×10^{-7}	$(une_{t-1}-une_{t-7})$	3.27×10^{-22}	0.010	2.49×10^{-4}	9.80×10^{-22}
$(e_{t-1}-e_{t-8})$	1.98×10^{-12}	0.227	1.23×10^{-9}	5.90×10^{-6}	$(une_{t-1}-une_{t-8})$	5.30×10^{-23}	4.45×10^{-5}	2.21×10^{-4}	3.59×10^{-19}
$(e_{t-1}-e_{t-12})$	5.48×10^{-11}	1.42×10^{-3}	6.73×10^{-4}	1.94×10^{-7}	$(une_{t-1}-une_{t-12})$	1.56×10^{-17}	7.04×10^{-4}	3.75×10^{-4}	1.33×10^{-14}
					$(une_{t-1}-une_{t-53})$	3.42×10^{-3}	0.046	0.217	9.63×10^{-3}
					$\widetilde{une}_t$	2.06×10^{-3}	0.571	4.37×10^{-4}	0.062

Note: The column headed s_t defines the candidate transition variable. Entries are are approximate p–values for the LM tests of nonlinearity (H_0'), and for the sequence of tests defined in (17)–(19) for determining whether the transition function is likely an exponential or a logistic.

Table 4: Single–Equation Model Assessment and Diagnostic Test Results.

	Southeast U.S. Price		Eastern Canada Price		Exchange Rate		Unemployment Rate	
Measure	Linear	STR Model	Linear	STR Model	Linear	STR Model	Linear	STR Model
Type	–	LSTR	–	LSTR	–	ESTR	–	LSTR
s_t	–	$\widetilde{une}_t$	–	$\widetilde{une}_t$	–	Δe_{t-1}	–	$une_{t-1} - une_{t-4}$
R^2	0.316	0.366	0.276	0.330	0.021	0.059	0.410	0.584
$\hat{\sigma}_v$	0.052	0.050	0.053	0.051	0.013	0.013	0.095	0.080
$\hat{\sigma}_{v,\mathrm{NL}}/\hat{\sigma}_{v,\mathrm{L}}$	–	0.955	–	0.962	–	0.981	–	0.840
AIC	-3.063	-3.096	-2.991	-3.026	-5.800	-5.806	-1.811	-2.086
AR(4)	0.538	0.893	0.287	0.077	0.235	0.965	0.676	0.208
AR(6)	0.758	0.860	0.420	0.214	0.389	0.506	0.747	0.006
AR(12)	0.699	0.523	0.848	0.561	0.206	0.708	0.437	0.030
ARCH(4)	7.30×10^{-7}	6.43×10^{-5}	6.43×10^{-5}	3.20×10^{-3}	4.01×10^{-26}	1.46×10^{-27}	3.13×10^{-12}	2.05×10^{-14}
ARCH(6)	3.77×10^{-6}	1.81×10^{-4}	2.64×10^{-9}	4.37×10^{-5}	2.04×10^{-28}	6.22×10^{-29}	3.54×10^{-19}	2.50×10^{-13}
SK	-0.164	-0.034	0.026	0.227	-1.024	-1.034	0.800	0.481
EK	2.332	2.054	2.533	2.153	6.855	7.406	2.937	3.157
LJB	140.25	106.77	162.38	122.41	1294.40	1495.43	282.88	275.52

Note: The effective sample size, T, is 607. LSTR denotes logistic smooth transition and ESTR exponential smooth transition. s_t indicates the transition variable used. R^2 is the unadjusted R^2 and $\hat{\sigma}_v$ is the residual standard error. $\hat{\sigma}_{v,\mathrm{NL}}/\hat{\sigma}_{v,\mathrm{L}}$ is the ratio of the respective standard error from the nonlinear model relative to the linear model. SK is skewness, EK is excess kurtosis, and LJB is the Lomnicki–Jarque–Bera test of normality of the residuals (critical value from the χ^2_2 distribution is 13.82 at the 0.001 significance level). AR(j), $j = 4, 6, 12$, is the p–value from an F–version of the LM test for remaining autocorrelation up to lag j. Entries for ARCH(j), $j = 4, 6$ are similarly defined for ARCH errors up to lag j.

Table 5: STVECM Model Estimates.

Panel A: Southeast U.S. Price, $y_{1t} = ln(p_{it})$

$$\Delta y_{1t} = [\underset{0.0018}{-0.0015} + \underset{(0.019)}{0.151}\,\Delta y_{1t-1} - \underset{(0.020)}{0.195}\,\Delta y_{1t-2} - \underset{(0.021)}{0.051}\,\Delta y_{1t-3} - \underset{(0.019)}{0.028}\,\Delta y_{1t-4} + \underset{(0.021)}{0.302}\,\Delta y_{2t-1} - \underset{(0.017)}{0.048}\,\Delta y_{2t-2} - \underset{(0.020)}{0.137}\,\Delta y_{2t-3} - \underset{(0.019)}{0.055}\,\Delta y_{2t-4}$$

$$- \underset{(0.022)}{0.076}\,\hat{\varepsilon}_{t-1}] \times [1 - G_1(s_{1t}, \gamma_1, c_1)] + [-\underset{(0.0031)}{0.0026} + \underset{(0.031)}{0.544}\,\Delta y_{1t-1} - \underset{(0.045)}{0.731}\,\Delta y_{1t-2} - \underset{(0.029)}{0.251}\,\Delta y_{1t-3} - \underset{(0.027)}{0.238}\,\Delta y_{1t-4} + \underset{(0.033)}{0.558}\,\Delta y_{2t-1} + \underset{(0.037)}{0.388}\,\Delta y_{2t-2}$$

$$+ \underset{(0.046)}{0.577}\,y_{2t-3} - \underset{(0.037)}{0.166}\,y_{2t-4} + \underset{(0.055)}{0.098}\,\hat{\varepsilon}_{t-1}] \times G_1(s_{1t}; \gamma_1, c_1) + \hat{v}_{1t}; \quad G_1(s_{1t}; \gamma_1, c_1) \;=\; [1 + \exp\{-\underset{(0.532)}{1.486}(s_{1t} - \underset{(0.014)}{2.908})/\hat{\sigma}_{s_{1t}}\}]^{-1},\ \mathrm{R}^2 = 0.352.$$

Panel B: Eastern Canada Price, $y_{2t} = ln(p_{xt})$

$$\Delta y_{2t} = [\underset{0.0019}{-0.0015} + \underset{(0.019)}{0.178}\,\Delta y_{1t-1} - \underset{(0.021)}{0.147}\,\Delta y_{1t-2} + \underset{(0.024)}{0.018}\,\Delta y_{1t-3} - \underset{(0.024)}{0.041}\,\Delta y_{1t-4} + \underset{(0.020)}{0.356}\,\Delta y_{2t-1} - \underset{(0.019)}{0.100}\,\Delta y_{2t-2} + \underset{(0.023)}{0.096}\,\Delta y_{2t-3} - \underset{(0.021)}{0.030}\,\Delta y_{2t-4}$$

$$- \underset{(0.024)}{0.032}\,\hat{\varepsilon}_{t-1}] \times [1 - G_2(s_{2t}, \gamma_2, c_2)] + [\underset{(0.0038)}{0.0067} + \underset{(0.031)}{0.311}\,\Delta y_{1t-1} - \underset{(0.037)}{0.791}\,\Delta y_{1t-2} + \underset{(0.040)}{0.282}\,\Delta y_{1t-3} - \underset{(0.032)}{0.189}\,\Delta y_{1t-4} + \underset{(0.047)}{0.608}\,\Delta y_{2t-1} + \underset{(0.026)}{0.157}\,\Delta y_{2t-2}$$

$$+ \underset{(0.037)}{0.422}\,y_{2t-3} - \underset{(0.029)}{0.283}\,y_{2t-4} + \underset{(0.060)}{0.178}\,\hat{\varepsilon}_{t-1}] \times G_2(s_{2t}; \gamma_2, c_2) + \hat{v}_{2t}; \quad G_2(s_{2t}; \gamma_2, c_2) \;=\; [1 + \exp\{-\underset{(29.930)}{448.009}(s_{2t} - \underset{(0.003)}{2.764})/\hat{\sigma}_{s_{2t}}\}]^{-1},\ \mathrm{R}^2 = 0.322.$$

Panel C: Exchange Rate, $y_{3t} = ln(e_t)$

$$\Delta y_{3t} = [-\underset{(0.003)}{0.023} + \underset{(0.158)}{1.253}\,\Delta y_{3t-1} - \underset{(0.252)}{0.537}\,\Delta y_{3t-2} + \underset{(0.305)}{0.696}\,\Delta y_{3t-3} + \underset{(0.257)}{0.086}\,\Delta y_{3t-4} + \underset{(0.285)}{0.433}\,\Delta y_{3t-5} - \underset{(0.253)}{0.192}\,\Delta y_{3t-6}] \times [1 - G_3(s_{3t}, \gamma_3, c_3)]$$

$$+[-\underset{(0.00057)}{0.00068} - \underset{(0.027)}{0.039}\,\Delta y_{3t-1} + \underset{(0.020)}{0.002}\,\Delta y_{3t-2} - \underset{(0.027)}{0.073}\,\Delta y_{3t-3} - \underset{(0.032)}{0.044}\,\Delta y_{3t-4} - \underset{(0.033)}{0.057}\,\Delta y_{3t-5} - \underset{(0.028)}{0.130}\,\Delta y_{3t-6}] \times G_3(s_{3t}, \gamma_3, c_3) + \hat{v}_{3t},$$

$$G_3(s_{3t}; \gamma_3, c_3) \;=\; 1 - \exp\{-\underset{(1.262)}{7.703}(s_{3t} - \underset{(0.001)}{0.022})^2/\hat{\sigma}^2_{s_{3t}}\},\ \mathrm{R}^2 = 0.050.$$

Table 5: Continued.

Panel D: Unemployment Rate, $y_{4t} = une_t$

$$\begin{aligned}
\Delta y_{4t} = [&- \underset{(0.007)}{0.042} + \underset{0.015}{0.060} \sin(2\pi t/52) + \underset{0.006}{0.042} \cos(2\pi t/52) + \underset{0.012}{0.059} \sin(2\pi t/26) - \underset{0.013}{0.113} \cos(2\pi t/26) - \underset{0.010}{0.005} \sin(2\pi t/13) + \underset{0.008}{0.041} \cos(2\pi t/13) \\
&- \underset{(0.097)}{0.829}\, \Delta y_{4t-1} - \underset{(0.081)}{0.104}\, \Delta y_{4t-2} - \underset{(0.094)}{0.049}\, \Delta y_{4t-3} - \underset{(0.062)}{0.0046}\, \Delta y_{4t-4} + \underset{(0.055)}{0.135}\, \Delta y_{4t-5} + \underset{(0.059)}{0.111}\, \Delta y_{4t-6} + \underset{(0.039)}{0.142}\, \Delta y_{4t-7} + \underset{(0.058)}{0.239}\, \Delta y_{4t-8} \\
&+ \underset{(0.038)}{0.141}\, \Delta y_{4t-9} + \underset{(0.044)}{0.052}\, \Delta y_{4t-10} + \underset{(0.043)}{0.021}\, \Delta y_{4t-11} - \underset{(0.002)}{0.015}\, y_{4t-1}] \times [1 - G_4(s_{4t}, \gamma_4, c_4)] + [- \underset{(0.003)}{0.059} - \underset{0.006}{0.012} \sin(2\pi t/52) + \underset{0.011}{0.128} \cos(2\pi t/52) \\
&+ \underset{0.008}{0.027} \sin(2\pi t/26) + \underset{0.008}{0.055} \cos(2\pi t/26) + \underset{0.009}{0.057} \sin(2\pi t/13) + \underset{0.009}{0.115} \cos(2\pi t/13) - \underset{(0.045)}{0.382}\, \Delta y_{4t-1} - \underset{(0.035)}{0.427}\, \Delta y_{4t-2} - \underset{(0.036)}{0.362}\, \Delta y_{4t-3} \\
&+ \underset{(0.029)}{0.273}\, \Delta y_{4t-4} + \underset{(0.036)}{0.866}\, \Delta y_{4t-5} + \underset{(0.028)}{0.453}\, \Delta y_{4t-6} + \underset{(0.044)}{0.330}\, \Delta y_{4t-7} + \underset{(0.037)}{0.424}\, \Delta y_{4t-8} + \underset{(0.036)}{0.295}\, \Delta y_{4t-9} - \underset{(0.036)}{0.115}\, \Delta y_{4t-10} + \underset{(0.041)}{0.374}\, \Delta y_{4t-11} \\
&- \underset{(0.001)}{0.012}\, y_{4t-1}] \times G_4(s_{4t}, \gamma_4, c_4) + \hat{v}_{4t}; \quad G_4(s_{4t}; \gamma_4, c_4) \;=\; [1 + \exp\{- \underset{(0.365)}{3.096}(s_{4t} - \underset{(0.019)}{0.149})/\hat{\sigma}_{s_{4t}}\}]^{-1},\ \mathrm{R}^2 = 0.583.
\end{aligned}$$

Panel E: Model Summary Statistics

$$\ln L = 4703.848,\ \tilde{\mathrm{R}}^2 = 0.780,\ AIC_{NL} = -26.511,\ SBC_{NL} = -20.448,\ AIC_L = -26.180,\ SBC_L = -25.802,\ \det\left(\hat{\Sigma}_{NL}\right)\Big/\det\left(\hat{\Sigma}_L\right) = 0.603,$$

$$\hat{\sigma}_1^2 = \underset{(4.50\times10^{-5})}{2.50 \times 10^{-3}},\ \hat{\sigma}_2^2 = \underset{(4.12\times10^{-5})}{2.65 \times 10^{-3}},\ \hat{\sigma}_3^2 = \underset{(5.63\times10^{-6})}{1.68 \times 10^{-4}},\ \hat{\sigma}_4^2 = \underset{(7.33\times10^{-4})}{6.36 \times 10^{-3}},\ \hat{\sigma}_{12} = \underset{(2.14\times10^{-5})}{2.14 \times 10^{-3}},\ \hat{\rho}_{12} = 0.832$$

Note: Asymptotic heteroskedasticity robust standard errors are given below parameter estimates in parentheses; R^2 is the squared correlation between actual and fitted values for each equation; $\hat{v}_{jt}$ denotes the j^{th} equation's residual at time t, $j = 1, ..., 4$; $\tilde{\mathrm{R}}^2$ denotes the likelihood system R^2 as defined by Magee (1990); AIC is the system Akaike information criterion and SBC denotes the system Schwarz's Bayesian Criterion. A subscripted L refers to the linear model and NL a nonlinear model. As well, $\det\left(\hat{\Sigma}_{NL}\right)\Big/\det\left(\hat{\Sigma}_L\right)$ denotes the ratio of the determinant of the covariance matrix for the STVECM relative to the VECM. $\hat{\sigma}_j^2$ denotes the estimated variance for equation j, $\hat{\sigma}_{12}$ is the estimated covariance term for the residuals between p_{it} and p_{xt}, and $\hat{\rho}_12$ is the corresponding correlation coefficient.

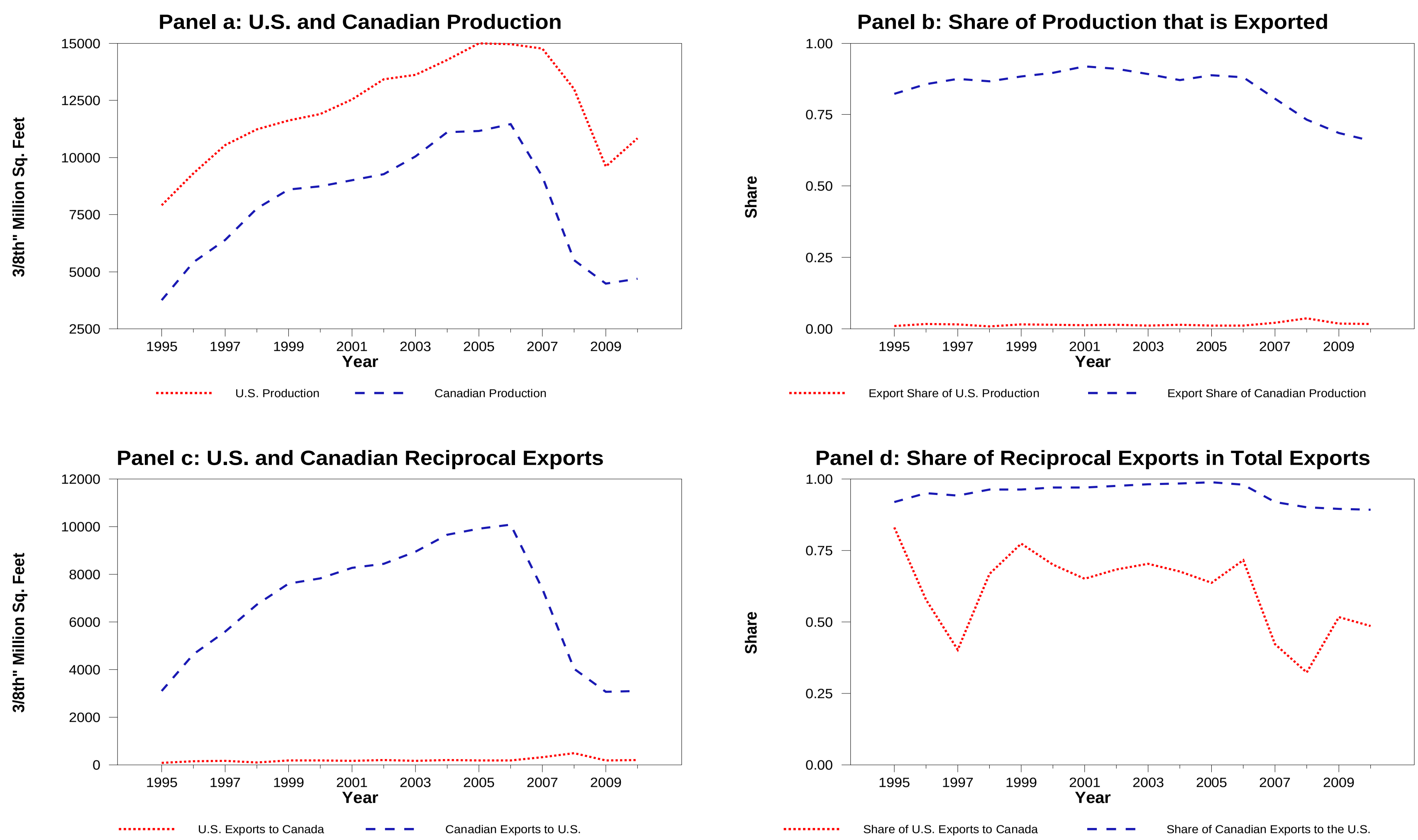

Figure 1: Production and Export Trends for 3/8th" OSB in the United States and Canada, 1995–2010. The data were obtained from annual reports of the Structural Board Association (SBA).

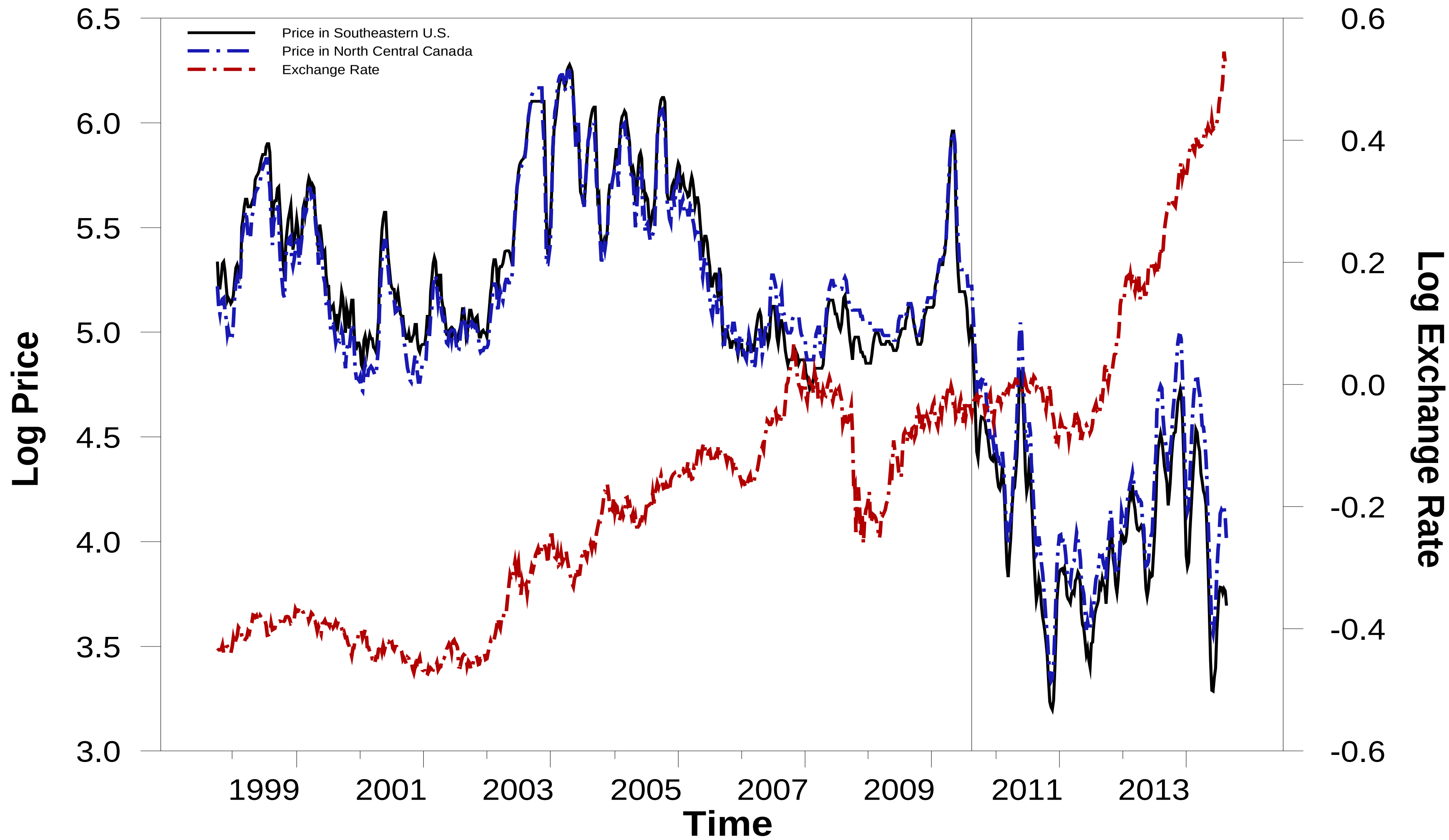

Figure 2: Natural Logarithms of OSB Prices in Northeastern Canada and Southeastern U.S. along with the U.S.–Canadian Dollar Nominal Exchange Rate, Actual and Simulated, 1998–2014.

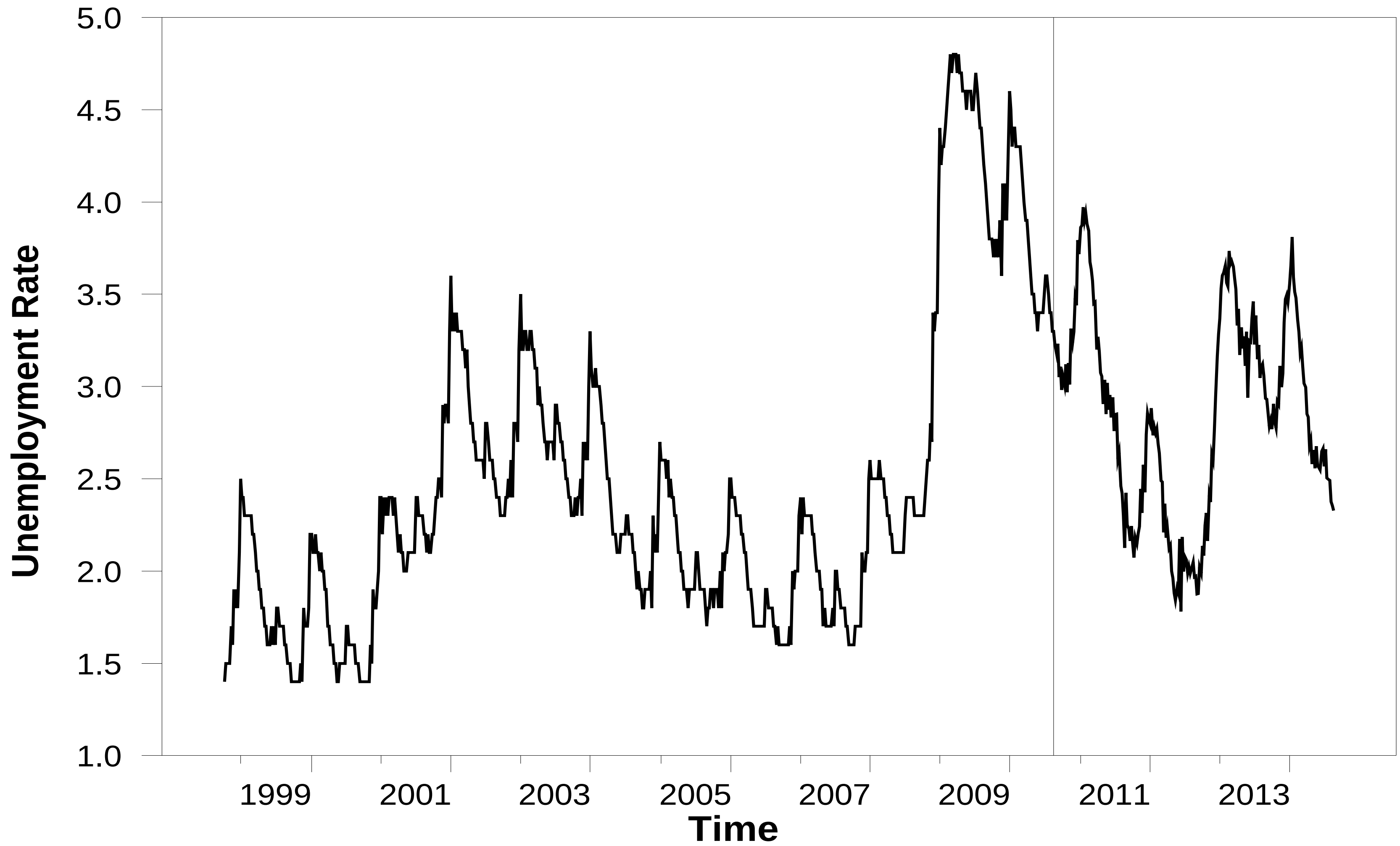

Figure 3: Weekly U.S. Unemployment Claims, Actual and Simulated, 1998–2014.

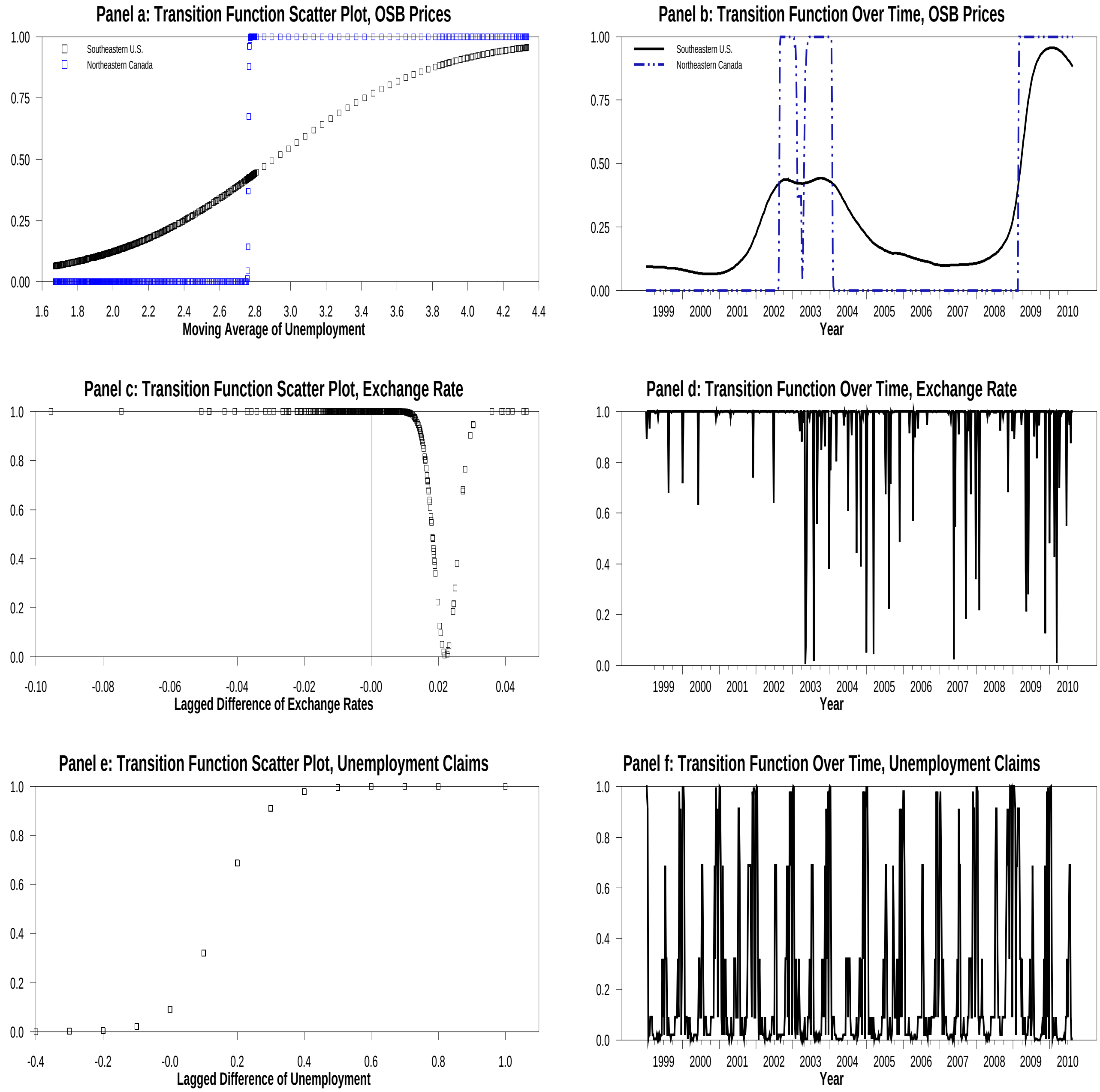

Figure 4: Transition Functions for the Estimated STVECM Model. Panels in the left–hand column show estimated transition functions plotted against corresponding transition variables. Panels in the right–hand column show the estimated transition function values over time.

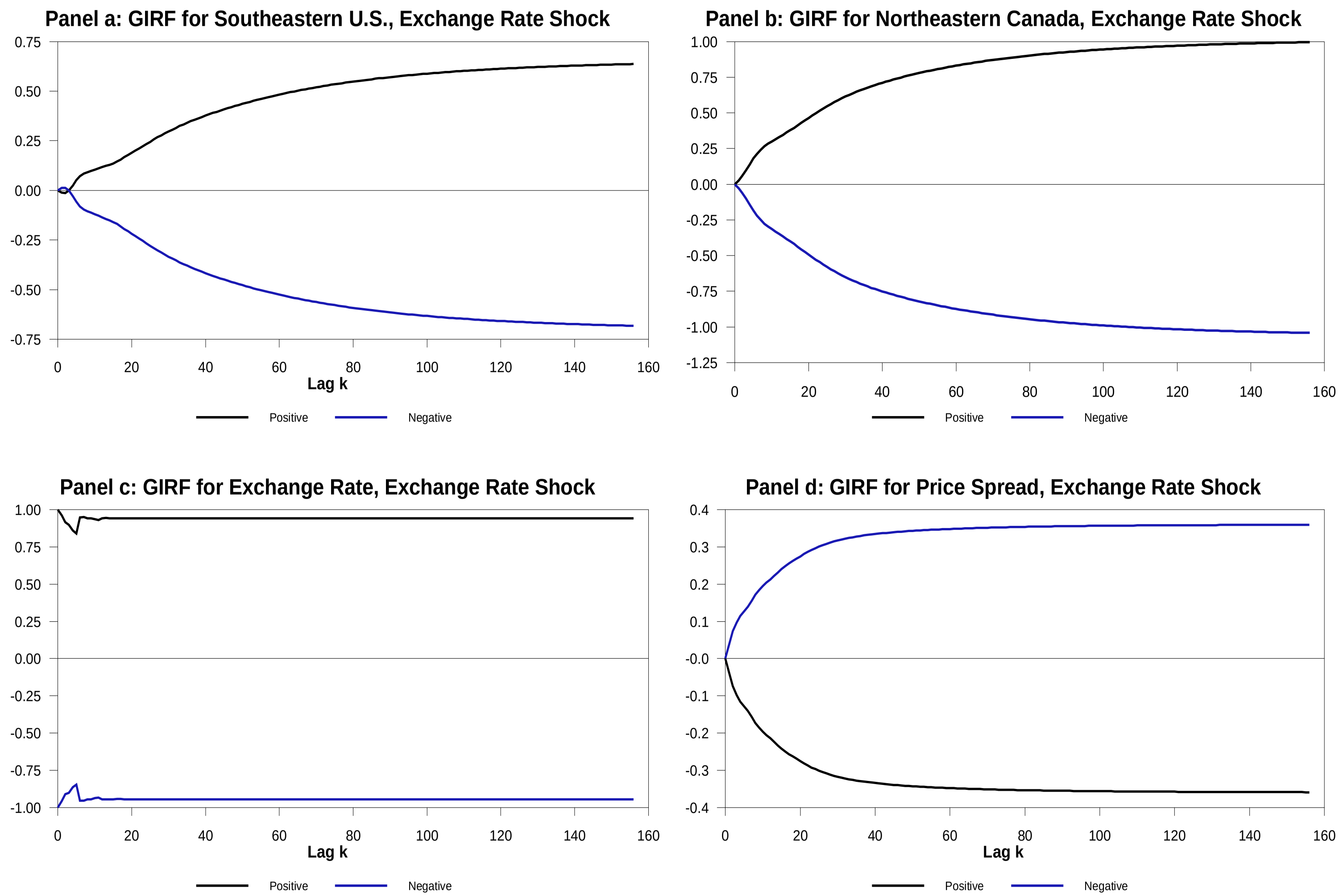

Figure 5: Unconditional Generalized Impulse Response Functions for Unit Shocks to U.S.–Canadian Dollar Exchange Rate.

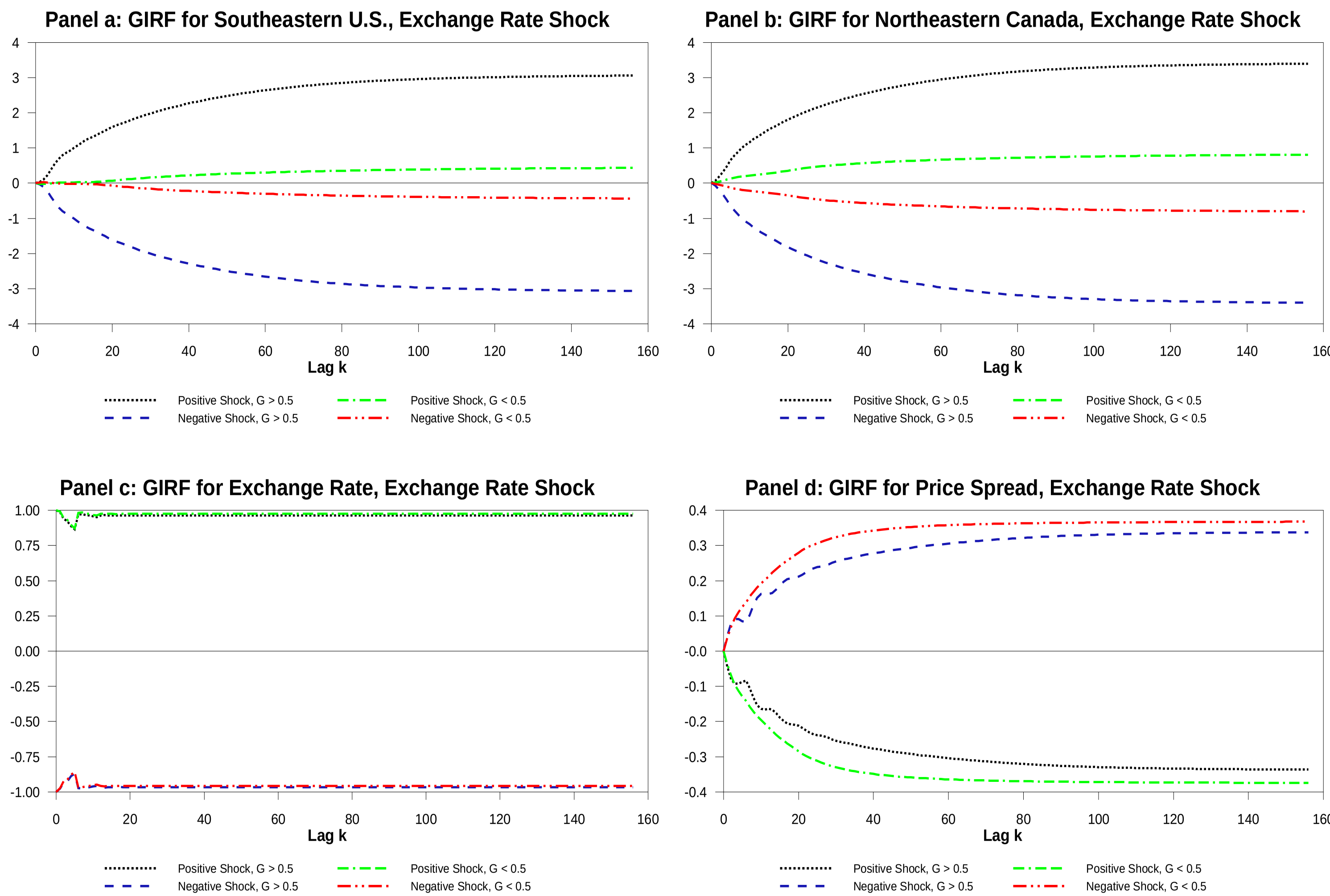

Figure 6: Generalized Impulse Response Functions for Unit Shocks to U.S.–Canadian Dollar Exchange Rate Conditional on the Moving–Average of Weekly U.S. Unemployment Rates being Greater than (Less than) 2.91–Percent at Horizon $n = 0$.

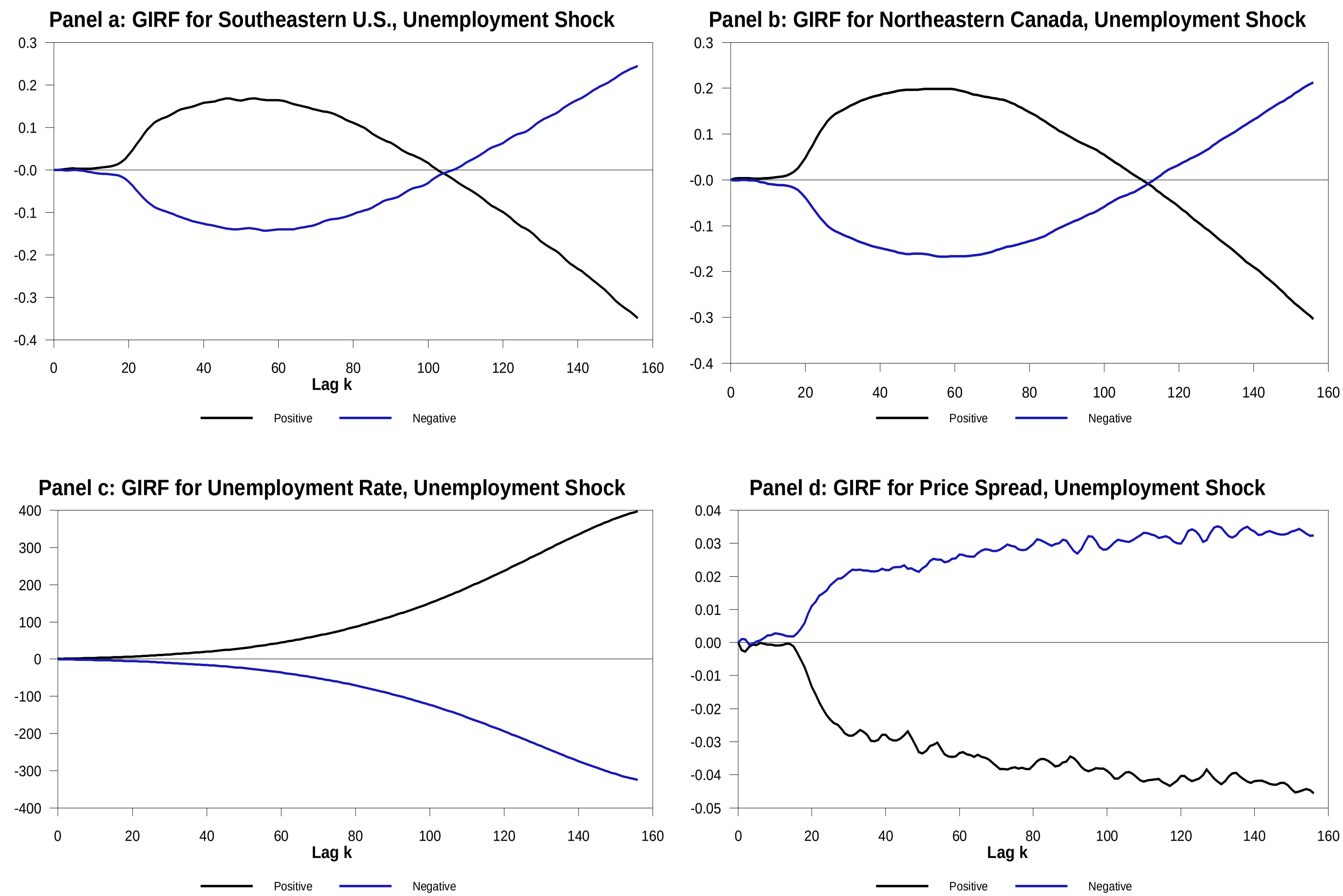

Figure 7: Unconditional Generalized Impulse Response Functions for a One– Standard–Deviation Shock to the U.S. Weekly Unemployment Rate.

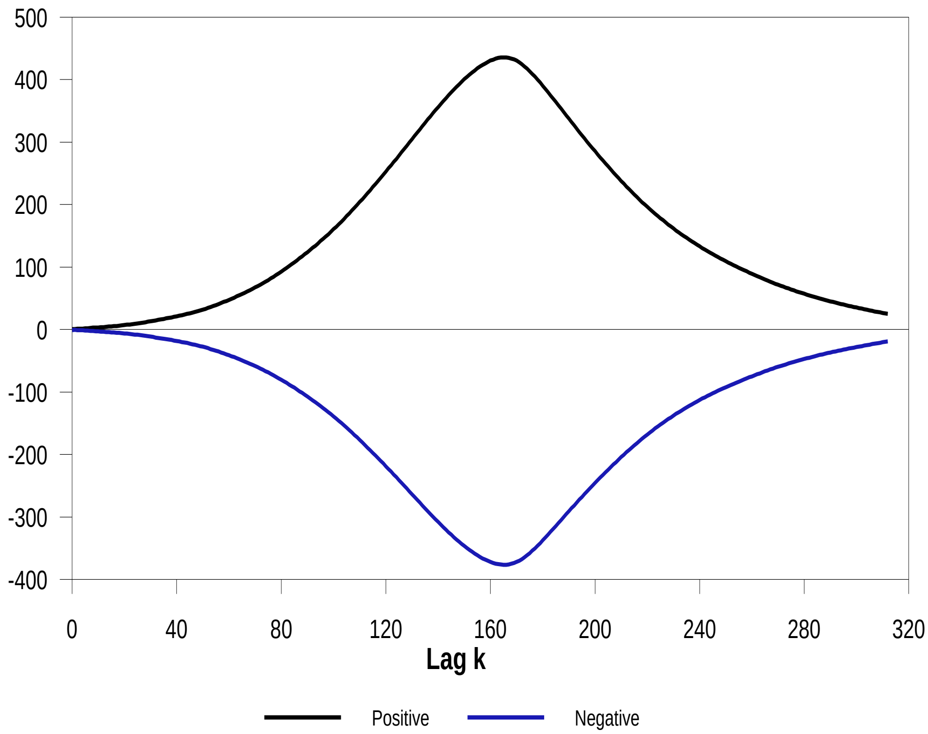

Figure 8: Unconditional Generalized Impulse Response Functions for a One–Standard–Deviation Shock to Unemployment Over a Six–Year Horizon.

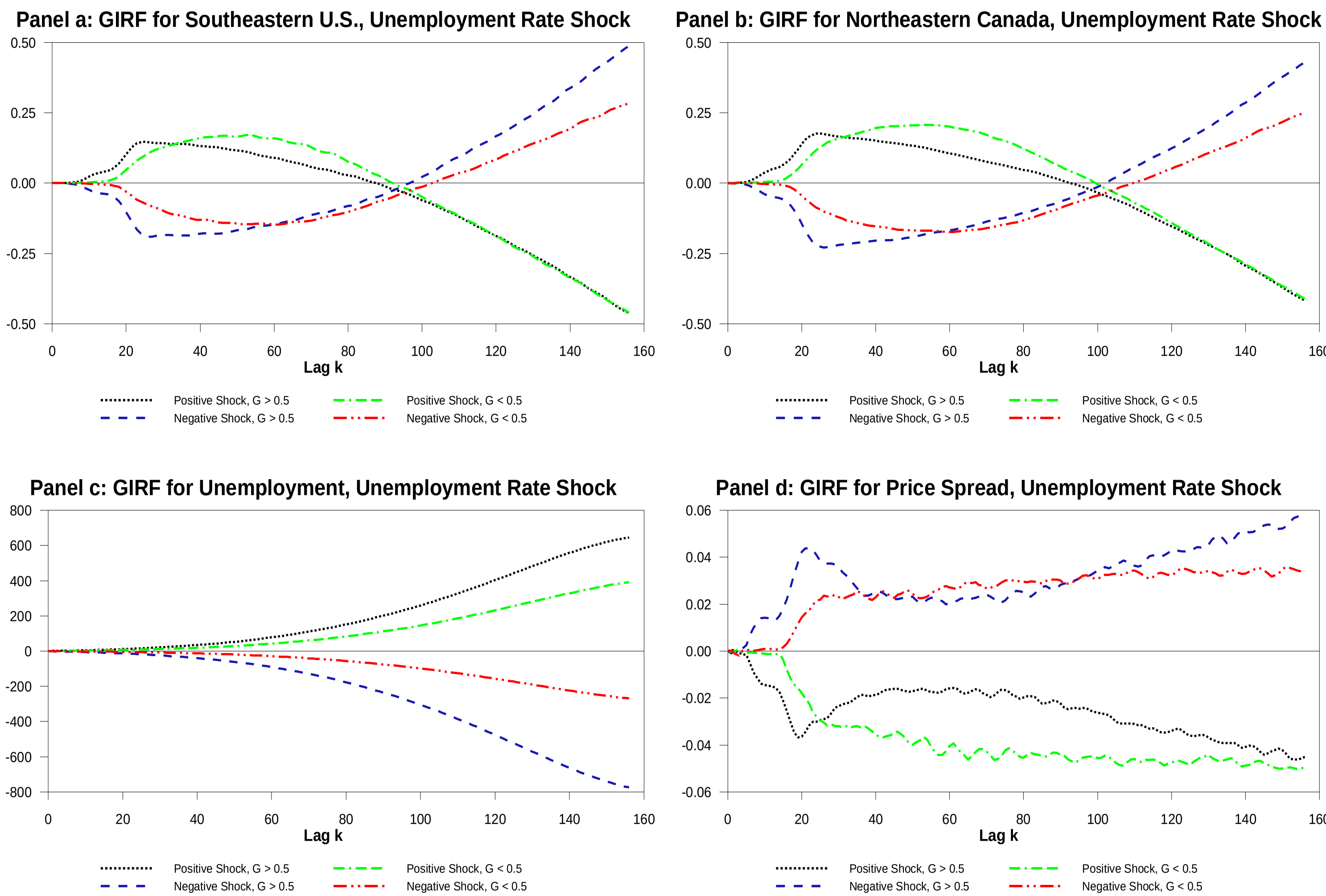

Figure 9: Generalized Impulse Response Functions Generalized Impulse Response Functions for a One– Standard–Deviation Shock to the U.S. Weekly Unemployment Rate Conditional on $G_4(s_{t4}) \geq$ (respectively $<$) 0.5.